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Essays on Purchasing Power Parity Puzzle

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ACADEMIC DISSERTATION

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List of Essays

1. Models and relations in purchasing power parity literature.

2. Long-run deviations from the purchasing power parity between the German mark and the U.S. dollar: Oil price - the missing link?

3. The U.S. dollar real exchange rate. A real options' approach.

4. Purchasing power parity puzzle: A sudden nonlinear perspective.

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MODELS AND RELATIONS
IN PURCHASING POWER PARITY
LITERATURE

1. INTRODUCTION

"The purchasing power parities represent the true equilibrium of the exchange, and it is of great practical values to know these parities. It is in fact to them we have to refer when we wish to get an idea of the real value of currencies whose exchange are subject to arbitrary and sometimes wild fluctuations" Gustav Cassel

The fundamental idea of the purchasing power parity (hereafter PPP) condition is that the prices of goods should tend to equal one another when expressed in a common currency. Thus, the law of one price is a crucial building block of PPP. The basic argument for why the law of one price should hold is based on the idea of frictionless goods market arbitrage and perfect substitutability between goods across different regions. In the strongest version absolute PPP states that, in the absence of transportation and other transaction costs, competitive markets will equalize the prices of identical goods in two countries when the prices are expressed in the same currency.[1]

As a basis for the international comparison of income and expenditure, PPP based on an overall price index established common ground for cross-country comparison by linking the currencies of different countries to the same base. In this case, PPP is superimposed as an *a priori* condition to convert a country's income and expenditure in local currency to a common unit (Summers and Heston, 1991, p.329). However, due to problems in specifying comparable price indices in two countries, the majority of the empirical literature tries to verify the relative version of PPP. Relative PPP states the weaker condition that the exchange rate will be proportional to the ratio of money price levels between countries, i.e. the relative purchasing power of national currencies.

Time series testing of the PPP hypothesis, typically defined with respect to a general or overall price level, use a simple estimating equation of the following form

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[1] For example, the famous "Big Mac Index" measures the degree of price equalization across countries for McDonald's hamburgers.

$$s_t = \beta(p_t - p_t^*) + \varepsilon_t, \tag{1.1}$$

where s_t is the log of the nominal exchange rate, p_t denotes the log of the domestic price level and p_t^* is the log of the foreign price level. If the restriction $\beta = 1$ is imposed, the strong form of PPP, the residual is constructed rather than estimated using Equation (1.1). The residual is then termed the real exchange rate. Thus, the real exchange rate may be viewed as a measure of the nominal exchange rate deviation from the fundamental PPP equilibrium.

The aim of this essay is to introduce the progress made by profession in understanding real exchange rate behavior. We discuss both theoretical and empirical aspects of real exchange rate behavior. Finally, we shortly discuss the possible contributions in this dissertation.

2. HISTORICAL BACKGROUND OF THE PURCHASING POWER PARITY HYPOTHESIS

The purchasing power parity hypothesis is one of the oldest topics in international economics. Although the term "purchasing power parity" was coined as recently as 85 years ago by the Swedish economist Gustav Cassel (Cassel, 1918), it has a much longer history in economics. Indeed, it has no doubt been around as long as currencies have been exchanged, but it was probably first articulated by the scholars of the Salamanca school in sixteenth century Spain (Officer, 1982). The rise in interest in the purchasing power concept at that particular time was by no coincidence. The prohibition of usury by the Catholic Church forced lenders to justify interest payments. If lending in foreign currency, lenders could justify interest payments by reference to movements in purchasing power. De Molina (1601), cited in Grice-Hutchison (1993, p.165), summarized the discussion on purchasing power during the course of the previous century. He wrote: "Other things equal, whenever money is most abundant, there it will

be least valuable for the purpose of buying goods and comparing things other than money…We see that money is far less valuable in the New World than it is in Spain."

In the eighteenth century England adopted paper money and subsequently London became the world's principal financial center. During this time period, interest in exchange rate theory increased in England. The English philosopher David Hume first stated the purchasing power hypothesis more formally in 1752 and John Wheatley explained currency fluctuations using the quantity theory of money coupled with the purchasing power theorem.

The nineteenth century was a period of considerable international economic integration and, in particularly in the United States, a period of considerable economic growth. Britain adopted gold standard in 1821 and retain this regime, together with many other countries, right up to the outbreak of World War I. Intense international economic and financial integration provided a stable environment for international financial markets. Indeed, PPP worked very well under the classic gold standard before 1914, as noted by McCloskey and Zecher (1984). After high wartime inflation, adjusting exchange rates to be consistent with PPP was quickly seen as a macroeconomic problem (Taylor, 1996, p.3). Rogoff (1996) emphasizes that the modern origins of PPP can be traced to the debate on how to restore the world financial system after the collapse of the gold standard during World War I.

While PPP theory goes back to Hume and Wheatley, the modern discussion on the purchasing power parity relation was intensified by the articles of the Swedish economist Gustav Cassel in the late 1910's. He was motivated by the vast dispersion in national price levels driven by wartime inflation in various countries. When World War I ended countries faced the very real problem of deciding how to reset exchange rates with minimal disruption to prices and government finances. Cassel's perception of the severity of parity dislocations after War World I was remarkable. Cassel's PPP calculations

played an especially important role in the debate over Britain's much-criticized decision to try to restore its prewar mint parity with the dollar in 1925 (Rogoff, 1996).[2]

Cassel (1922) criticized the use of the term "high" and "low" exchange rates and argued that in reality PPP presents an indifferent equilibrium of the exchange in the sense that it does not affect international trade either way. However, it is important to note that Cassel discusses several limitations of PPP throughout of his writings (Holmes, 1967). Cassel's view was that a short term speculation or higher expected inflation in the home country than abroad may, among other things, move the exchange rate away from PPP.[3] Thus, Cassel's arguments are partly based on a sophisticated theory of price expectations, as noted by Officer (1976).

Although it was widely accepted that commodity arbitrage is essential for PPP, it was not self-evident in those days whether PPP refer to traded commodities only or a broader basket of goods. This debate goes back as far as 1821, when Ricardo stated: "The exchange is never ascertain by estimating the comparative value of money in corn, cloth or any commodity whatever but by estimating the value of the currency of one country, in the currency of another". Since then the choice of the appropriate price index to be used in implementing PPP has been the object of a long debate. Heckscher (1916) pointed out that PPP is based on commodity arbitrage, i.e. the basis for PPP is the law of one price. Hence, purchasing power refers to purchasing power on tradables. Furthermore, Heckscher argues that international transaction costs should create some scope for deviations from PPP. Cassel (1922) viewed PPP as the equilibrium exchange rate based on general price levels of the countries, representing all goods and services available for purchase. Cassel's view was supported by Keynes (1930). Keynes pointed out that PPP calculated from traded goods prices alone is close to truism.

[2] Note also Keynes's pamphlet: The Economic Consequences of Mr. Churchill, where Keynes criticizes the then Finance minister for returning the U.K. to the gold standard at pre-war parity.
[3] Pigou (1920) extended the sources of limitations of PPP and introduced a concept of third-degree price discrimination to refer to integrated and segmented goods markets.

There was also an extensive debate on the different nature of PPP. The Casselian PPP approach views the exchange rate as the determined variable and price levels as causal variables, whereas others noted that there are also chains of causation running from exchange rates to prices (e.g. Keynes, 1923). Subsequently Einzig (1935) also pointed out that changes in exchange rates produce changes in relative prices, which is a contradiction to the Casselian PPP theory of exchange. Einzig's observation corresponds to the notation during the First World War that in a system of flexible exchange rates appreciation of a country's currency leads to a decrease in the general price level because of their impact on domestic activity.

In 1944 in Bretton Woods, New Hampshire, the major Allied Powers drew up plans for an international monetary system of fixed exchange rates in which the United States dollar would effectively be the reserve currency (Lothian, 2001, p.10). Exchange rate equilibration under a combined gold and dollar standard was potentially feasible, but often blocked by political fears of the cost of adjustment in the countries which needed to reflate (Eichengreen, 1992). Finally in early 1973 the regime broke down and the floating exchange rates took the place of the Bretton Woods system.

Cassel's theory on PPP accepts the fact that there are nontraded goods but notes that the prices of traded and nontraded goods are closely related through various links (Officer, 1976, p.8). Harrod (1939) and later Balassa (1964) and also Samuelson (1964), examined more carefully the consequences of nontradable goods for the theory of PPP. They were motivated by the empirical regularity that wealthy countries have higher price levels than poor countries. They concluded that the primary effect of traded goods productivity growth is increased wages in the traded goods sector. Since labor tends in the long run to be mobile across sectors, wage increases in the traded goods sector push up wages in the nontraded sector. Since there is slower productivity growth in the nontraded sector, an increase in wages in the traded sector is passed along into higher prices of nontraded goods.

The open economy monetarist model in the 1960s and 1970s relied on the assumption that at least relative PPP holds. Indeed, prior to the recent float, the professional consensus appeared to support the existence of a varying but fairly stable real exchange rate (Sarno and Taylor, 1997). The historical success or failure of PPP was seen as intimately tied to the mobility of global financial capital. In the mid to late 1970s, in the light of very high variability of the real exchange rates after the major exchange rates were allowed to float, this position was largely abandoned. The proposition that the even real exchange rates are volatile when nominal exchange rates are allowed to float freely has become something of a stylized fact in the international real exchange rate literature (Froot and Rogoff, 1995).

The reason for the problems in the Bretton Woods regime was the ease with which the shocks were transmitted internationally under pegged exchange rate arrangements. The reason for problems in the floating exchange regime seem to arise in the area of exchange rate behavior itself. Thus, despite the fact that there is a long history of PPP research, it is not surprising that PPP is one of the main tenets in the research of international economics even today.

3. THEORETICAL MODELS FOR PPP DEVIATIONS

Deviations from PPP are at the same time persistent and volatile. If the sources of PPP disturbances were real in nature, persistent deviations could be explained by real shocks. However, deviations are too volatile to be accounted for real shocks. The adjustment towards parity is also too slow to be explained by nominal rigidities. Thus, we have a purchasing power parity puzzle (Rogoff, 1996).

3.1 PERSISTENT DEVIATIONS FROM PPP

The Harrod-Balassa-Samuelson hypothesis is a cornerstone of the PPP literature, which tries to rationalize the existence of long-run deviations from PPP.[4] The Harrod-Balassa-Samuelson (hereafter HBS) hypothesis divides real exchange rate movements into two components. The first component of the hypothesis is the assumption that the relative price of non-tradables is proportional to the ratio of labor products and the second component is the assumption of PPP for traded goods. In this subsection we first analyze the HBS hypothesis more carefully to understand the sources of persistent deviations from PPP. In the following subsections we will discuss the assumption of PPP for traded goods more in detail.

Obviously, there are also several other important contributions based on real shocks in addition to HBS which may explain persistent deviations from PPP. The ratio of government spending to GDP is often seen to appreciate the real exchange rate. (e.g. De Gregorio *et al.* 1994). This is due to the fact that government consumption may fall more heavily on nontradable goods than on private consumption. Fluctuations in terms of trade may also affect the real exchange rate. From a theoretical perspective almost any correlation between the terms of trade and the real exchange rate can be easily rationalized. This depends on the channel through which a change in the terms of trade effect alters the real exchange rate, i.e. assumptions concerning intratemporal and intertemporal elasticity of substitution (Ostry, 1998). Movements in terms of trade also affect real exchange rate through the current account. Although the terms of trade and current account relations are ambiguous, the current accounts by themselves are likely to induce significant real exchange rate changes. This is because they lead to transfer of wealth across countries and home and foreign residents are likely to exhibit very different spending patterns, as noted by Krugman (1989).

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[4] Rogoff (1992) provides an alternative model for HBS based on intertemporal optimization of tradable goods consumption. The results are in sharp contrast to the predictions of the HBS model.

MacDonald and Ricci (2001) find, in turn, that an increase in productivity and in competitiveness of the distribution sector with respect to foreign countries leads to an appreciation of the real exchange rate. This effect is concurrent with the traditional Harrod-Balassa-Samuelson effect of productivity in the tradable and non-tradable sector.

3.1.1 THE HARROD-BALASSA-SAMUELSON HYPOTHESIS

The supply side is typically given by Cobb-Douglas production functions

$$Y_t^T = A_t^T (L_t^T)^\theta (K_t^T)^{1-\theta} \tag{3.1.}$$

$$Y_t^N = A_t^N (L_t^N)^\phi (K_t^N)^{1-\phi} \tag{3.2.}$$

where Y^T and Y^N are outputs of the traded and non-traded goods. L_t, K_t, and A_t represent labor input, capital input and stochastic productivity shock respectively. θ and ϕ are the labor shares in value added in traded and non-traded goods sectors. The marginal product of labor in tradable sector can be written as

$$\frac{\partial Y_t^T}{\partial L_t^T} = \theta \left(\frac{Y_t^T}{L_t^T} \right) = W \quad (3.3a) \qquad \text{or} \qquad \frac{\partial Y_t^T}{\partial L_t^T} = \theta A_t^T \left(\frac{K_t^T}{L_t^T} \right)^{1-\theta} = W \quad (3.3b)$$

The first equation (3.3a) shows that the marginal product of labor is proportional to the average product of labor with Cobb Douglas technologies. Perfect competition now implies that labor is paid the value of its marginal product (W). With perfect international mobility of the capital, profit maximization implies for capital that the capital-labor ratio in the traded sector is tied by the equation

$$\frac{\partial Y_t^T}{\partial K_t^T} = (1-\theta) A_t^T \left(\frac{K_t^T}{L_t^T} \right)^{-\theta} = R \tag{3.4}$$

where R is an international interest rate level. Taking tradables as numeraire price, we can write similar equations for non-tradables

$$\frac{\partial Y_t^N}{\partial L_t^N} = P_t^N \phi A_t^N \left(\frac{K_t^N}{L_t^N} \right)^{1-\phi} = W \tag{3.5}$$

Labor mobility across sectors guarantees that the nominal wage is equal in the two sectors. Thus, the price level is determined by the productivity differential between two sectors. Finally, capital-labor ratio in the non-tradable sector is determined by the equation

$$\frac{\partial Y_t^N}{\partial K_t^N} = P_t^N (1-\phi) A_t^N \left(\frac{K_t^N}{L_t^N} \right)^{-\phi} = R \tag{3.6}$$

Logarithmically differentiating Equations 3.3.b and 3.5 we can conclude that faster productivity growth in tradables than in non-tradables will push the price of non-tradables upward over time.

Let a "hat" above a variable denote logarithmic derivative $\hat{X} \equiv d \log X \equiv \frac{dX}{X}$ for any variable X restricted to assume positive values. We can write

$$P_t^N = \left(\frac{\phi}{\theta} \right) \hat{A}_t^T - \hat{A}_t^N \tag{3.7.}$$

Under the reasonable assumption that the labor share in non-tradables (ϕ) is higher than in tradables (θ), we obtain the expected result that faster productivity growth in tradables will push the price of nontradables upward over time. The conclusion is, *ceteris paribus*, that economies with a higher level of productivity in tradables will thus be characterized by higher wages and also by higher prices of non-tradables if productivity in non-tradable sector does not increase to the same extent, i.e. the economy will face a more appreciated real exchange rate.

Ultimately HBS requires only that the income share of labor is roughly constant and labor is mobile between sectors. These assumptions are realistic, especially in the long run. The PPP for tradables assumption, however, is under much debate. Next we discuss this issue in more detail.

3.2 INCOMPLETE PASS-THROUGH

Dornbusch (1987) pointed out that if demand curves have constant price elasticities in both foreign and domestic markets, a monopolistically competitive firm will follow a constant mark-up pricing rule, and the relative price of its product will remain constant as the exchange rate fluctuates even if markets are efficiently segmented. Dornbusch (1987) applies the industrial organization approach and shows that the extent of price adjustment depends on product sustainability, the relative number of domestic and foreign firms, and market structure. Dornbusch does not explain why prices are not changed as often as exchange rates move. Goldberg and Knetter (1997, p. 1270) conclude years later: "Although there is substantial variation across industries, in many cases half or more of the effect of an exchange rate change is offset by destination specific adjustment of markup over cost.".

A number of possible reasons can be evinced for the failure to find evidence of PPP for tradables. These include traditional forms of price stickiness and more modern ideas on local currency pricing (Devereux and Engel, 2000) as well as explanations based on price discrimination (Krugman, 1987), variable trade costs (Dumas, 1992) and sunk costs of arbitrage (Baldwin 1988).

3.2.1 STICKY PRICES

The high degree of correlation between movements in the nominal exchange rate and the real exchange rate is consistent with the hypothesis that prices of goods and services adjust sluggishly relative to asset prices, such as nominal exchange rates. If prices in goods markets are generally regarded as being sticky, volatility in nominal exchange rates is transferred into comparable real exchange rates. Thus, sticky prices are one explanation commonly evinced for real exchange rate fluctuations. The observed half-life persistence on the real exchange rate seems to be excessively high to rationalize by sticky prices. Price level movements do not begin to offset exchange rate swings on a monthly or even annual basis (Froot and Rogoff, 1995, p.1648). If nominal stickiness were really responsible for short-run PPP deviations one would expect substantial convergence to PPP over one to two years, as wages and prices adjust to a shock.

Chari *et al.* (2000) find that sticky prices can help replicate persistence in the data, but only if there is willingness to accept long-lived price contracts up to 3 years. It is generally thought that price-settings contracts are shorter than this in practice. It is important to note, however, that the view that price stickiness is important in explaining real exchange rate dynamics is difficult to identify since market frictions are difficult to quantify. If sticky prices are after all important in determining real exchange rates, shocks that induce delayed price responses should also play an important role in real exchange rate variations. Using this approach Ng (2003) finds that US sticky prices have been the main source of real dollar exchange rate variations since the collapse of the Bretton Woods agreement. However, real exchange adjustment to US sticky price shocks has been found to be a reasonable quick in Ng (2003), which indicates that they cannot be solely responsible for real exchange rate persistence.

Engel and Morley (2001) observe that the root of the PPP puzzle may lie in the possible different speeds of convergence for nominal exchange rates and prices. In contrast to standard rational expectations sticky-price models, which impose the same reversion of speed for nominal exchange rates and prices, they examine an empirical model that

allows those variables to adjust at different speeds. Their results show that while prices converge relatively fast, nominal exchange rates converge slowly.

3.2.2 PRICING TO MARKET

Since Krugman's (1987) article, studies on prices and exchange rates have focused intensively on the issues of markup adjustment.[5] Moving outside the competitive market paradigm, pricing to market (PTM) behavior gives rise to impediments to goods arbitrage. Thus, PTM effectively prevents traditional arbitrage forcing PPP. The main feature of this theory is that the same goods can be given a different price in different countries when oligopolistic firms are supplying them. When markets are segmented and the price elasticities of demand are not constant, a monopolistically competitive firm's optimal pricing behavior can drive a wedge between the common currency prices of the same goods destined to different markets.

Any perfectly competitive market is characterized by the condition that prices equal marginal costs. A perfectly competitive market implies an integrated market. A segmented market implies, instead, the existence of market power, as noted by Goldberg and Knetter (1997). Sources of international market segmentation are examined by Engel and Rogers (1996).[6] They use detailed CPI data for US and Canadian cities to study the effects of distance and the border on relative price volatility. Although both sources of relative price volatility are significant, the border effect is the dominant factor. They also find that relative price volatility is better explained by nominal exchange rate volatility than by measures of trade barriers. Bergin and Feenstra (2001) show that non-constant demand structure is an important condition for generating PTM behavior in price-setting firms and for helping staggered contracts to generate endogenous persistence. However, while certain specifications of the model have been shown to be able to generate a very high degree of persistence, these require implausible parameter values.

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[5] Krugman (1987) and Dornbusch (1987) used a partial equilibrium setting. PTM has been adopted to general equilibrium setting, for example, by Chari et. al (2000) and Betts and Devereux (2000).
[6] See also Parsley and Wei (1996).

Pricing to market behavior requires an imperfectly competitive market structure under which firms behave as price setters. Thus, it is quite conceivable that differences in market structure across industries play an important role in determining the persistence of deviations from PPP. Indeed, results based on disaggregated data in Cheung *et al*. (1999) show that differences in market structure significantly determine the rates at which deviations from sectoral PPP decay. Haskel and Wolf (2001) find that local distribution costs, local taxes and tariffs do not completely explain the price differences between different countries, leaving PTM resulting in varying markups.

Froot and Klemperer (1989) show that a model with consumer switching costs will lead exporters to respond differently to temporary and permanent changes in exchange rate. They examine the effects of temporary appreciation of the dollar focusing on dynamic demand side effects in an oligopolistic market. In their model temporary appreciation increases the value of current, relative to future, dollar profits expressed in foreign currency. When the dollar is temporarily high, foreign firms will find investments in market share less attractive, and will prefer instead to let their current profits increase.

Rogoff (1996) has cast doubt on PTM as an explanation for the persistence of real exchange rate especially if aggregate price indices are considered. PTM takes the ability to price discriminate to be absolute.[7] This may be the case for some goods, such as automobiles, where differences in national regulatory standards combined with the need for warranty service allow firms great leeway to price discriminate across countries. However, if we consider goods included in a tradable part of the CPI index, there is a substantial number of tradable goods which are homogenous in different countries.

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[7] Knetter (1993) finds that PTM is important for German and Japanese firms relative to US companies and it is strategy used a very broad range of goods.

3.2.3 LOCAL OR PRODUCER CURRENCY PRICING ?

The implications of pricing to market for the real exchange rate have been studied by comparing the behavior of the nominal exchange rate and prices in regimes with polar pricing rules. In the former, imports have been set in producer's currency as has been the traditional assumption in the Mundell-Fleming open economy macro models. Pass-through is complete when the response of import prices to exchange rate movements is one-for-one. In the standard Mundell-Fleming setup, the assumption of complete pass-through is related to the adjustment process of the current account to exchange rate movements.[8]

In the second system, import prices are set in consumer currencies, in line with the pricing to market literature. Persistent price differentials incorporate pricing to market by producers. Nominal price stickiness in prices denominated in the currency of the consumer, i.e. sticky local currency pricing is discussed in several papers (see e.g. Deveraux and Engel, 2000; Chari *et al*. 2000). Furthermore, as the elasticity of substitution rises, exporting firms become more concerned with maintaining their prices in line with domestic competitors. This leads to increased price rigidities in local currency terms. Thus, the change in the price of traded goods relative to domestic substitutes should be taken into account when measuring the parity between prices in the exporting and importing countries.

Obstfeld and Rogoff (2000a) have criticized the assumption of sticky local currency pricing on a number of grounds. In particular, they assert that invoicing in the importer's currency is not a widespread practice and that trade invoicing practices typically apply to contracts of 90 days or less. Thus, this type of price stickiness is too infrequent and brief to fully explain the degrees of persistence in relative price movements observed.[9]

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[8] Perfectly elastic export supply leads to Marshall-Lerner condition, i.e. a devaluation improves a country's balance of trade if the sum of the import and export demand elasticities exceeds one.

Finally, it is important to understand the difference between local currency pricing and PTM. If PTM is assumed, firms are able to adjust their own prices instantaneously when there are shifts in supply or demand, i.e. there is no fundamental price stickiness. Instead, in local currency pricing prices are completely sticky. In the flexible price setting, the currency in which the price is expressed is irrelevant, as noted by Engel (2003).

3.2.4 TRANSACTION COST MODELS

The sticky price explanation discussed above may explain the variability in real exchange rates since after the lag nominal exchange rate changes will translate one-for-one into real exchange rate changes. As discussed above, however, the models based on sticky prices and local currency pricing are insufficient on their own to explain the observed degree of real exchange rate persistency.

Dumas (1992) introduced real rigidities into the model in the form of international transaction costs between spatially separated markets. Goswami *et al*. (2002) also take account of a potential reduction in unit cost of distribution due to economies of scale. They show that this may help to understand the large short-term volatility observed in the real exchange rate. Aizenman (2000) focuses on time dependent transportation costs. The assumption in Aizenman (2000) is that the cost of delivering a good ordered ahead of time is lower than the cost of last minute delivery. It follows that in countries where terms of trade volatility is small, most imports are pre-bought, and the spot market for imports is inactive. Another implication of time dependent transportation costs is that higher financing costs would increase the cost of prepaying. This reduces the frequency of pricing to market thus increasing the tendency of the relative PPP to hold.

Financing costs may also have an effect on pricing behavior. Ahtiala and Orgler (1995) show that the optimal prices in the different currencies are equal at the spot exchange rate only by chance. By taking into consideration the impact of prices on sales in different

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[9] Obviously, manufactures are only one link in the supply chain and retailers are a more natural source of sticky local currency pricing. The results in Campbell and Lapham (2002) do not, however, support this assumption.

currencies, as well as all other relevant costs and risks, the exporter can optimally convert prices from one currency to another.

If there are frictions in international trading and these are time invariant, deviations from PPP should be constant over time. Pure transaction costs, however, are only a small proportion of traded goods prices. Rogoff (1996) offers a crude estimate of international shipping costs by comparing the Fob values with the Cif values.[10] This difference is estimated to be approximately 10 percent. Hummels (1999) estimates the average trade-weighted freight cost in the US in 1994 to be 3,8%. Therefore factors beyond pure transaction costs are needed to explain the deviations from the law of one price and PPP. By allowing a transaction cost which separates markets, however, it is possible to develop PTM models that can generate substantial persistence (Obstfeld and Rogoff, 2000b).

Almost all models of the real exchange rates that incorporate trade costs use Samuelson's iceberg formulation (e.g. Dumas, 1992; Coleman,1995; Obstfeld and Rogoff 2000b). Proportional transaction costs imply symmetric behavior of the real exchange rate according to whether it is above or below the equilibrium level. Indeed, Taylor *et al.* (2001, p.1021) state that:

"It is hard to think that of economic reasons why the speed of adjustment of the real exchange rate should vary according to whether the dollar is overvalued or undervalued, especially if one is thinking of goods arbitrage as ultimately driving the impetus toward the long run equilibrium and one is dealing with major dollar exchange rates against the currencies of other developed industrialized countries."

Dixit (1989a and b) shows using a real option theory that if firms face sunk cost of investment when breaking into foreign markets, the extent of pass-through will depend on the expected changes of exchange rate, i.e. the expected variance of the exchange rate is an important component of the determination of real exchange rates. This allows the

possibility that if risks are asymmetric then the adjustment path of the real exchange rate towards parity level is also asymmetric.[11]

Assuming market sunk costs of entry, sufficiently large real exchange rate shocks may alter domestic market structure and thereby induce hysteresis, as noted by Baldwin and Krugman (1989). O'Connel and Wei (1997) allow for fixed as well as proportional costs of arbitrage. This results in a two-threshold model where the real exchange rate is reset by arbitrage to an upper or lower inner threshold whenever it hits the corresponding outer threshold. Intuitively, arbitrage will be heavy once it is profitable enough to outweigh the initial fixed cost, but will stop short of returning the real rate to the PPP level because of the proportional arbitrage costs.

3.2.5 CURRENCY INVOICING AND PASS-THROUGH

There is abundant evidence showing that there has been a reduction in the pass-through of changes in exchange rates to consumer prices during last three decades (e.g. Goldberg and Knetter, 1997, McCarthy, 1999, Cagnon and Ihrig, 2001). There are two import factors that determine the extent of pass-through: the responsiveness of markups to competitive conditions and the degree of returns to scale in the production of imported goods. As an example consider the foreign firm which sets the price of a good exported to the United States as a constant markup over marginal cost. A complete pass-through occurs when returns to scale are constant. Thus, the change in market structure may explain decline in pass-through.

Taylor (2000) argues that the decline in pass-through is due to a reduction in the pricing power of firms. This, in turn, is caused by the low inflation environment achieved in many countries. Pass-through may also be endogenous to a county's monetary stability, i.e. countries with stable monetary policies would have their currencies chosen for transaction invoicing (Betts and Devereux, 2000; Devereux and Engel, 1999). Campa and

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[10] The Fob value is the value of world exports exclusive of transportation and insurance costs. The Cif value is the values of world imports inclusive of transport and insurance.
[11] For evidence for asymmetric adjustment, see Sollis *et al*. (2002).

Goldberg (2002), however, show that changes in inflation account for only a small fraction of the observed changes in pass-through elasticities. If the countries with very high inflation regime are considered, the relation between inflation and pass-through is clearer (De Grauwe and Grimaldi, 2002). However, De Grauwe and Grimaldi build their explanation on transaction costs.

3.3 MONETARY MODELS

Monetary models have also been proposed to explain real exchange rate fluctuations. For example, the celebrated Dornbusch (1976) overshooting model attributes the short term deviations from PPP due to stickiness in nominal prices. The empirical evidence for the overshooting model is rather weak (Faust and Rogers, 2003). Campbell and Clarida (1988) find that the real dollar exchange rate is so volatile and persistent that only a small fraction of the movement in the real exchange rate can be explained by the movements in the real interest rate differentials.[12] Thus, exchange rate changes are difficult to explain, at least over short horizons that match short-term interest rate horizons. The solution to this problem may emerge from information problems of market agents, as suggested in Faust and Rogers (2003).

Nakagava (2002) introduces threshold nonlinearity into a traditional real interest rate model to take account of a transaction cost-induced band of inaction for price adjustment. The model is able to establish a stronger link between real exchange rates and real interest rate differentials. Other explanations for the short term exchange rate volatility in the monetary models include financial factors such as changes in portfolio preferences and short term asset price bubbles, but such models cannot generate the observed slow convergence to PPP.[13]

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[12] Eichenbaum and Evans (1995) find evidence of a substantial link between monetary policy and exchange rates.

[13] A high degree of persistency in the bilateral real interest rate differential is capable of contributing to the persistence of the real exchange rate.

4. EMPIRICAL MODELS FOR PPP DEVIATIONS

4.1 PRICE INDICES

Typically, economists use a price index, like the CPI, to summarize the level of prices in each country. Because price indices are relative to base year, they do not give any indication OF how large absolute PPP deviations were for the base year. An important problem with the notation of PPP is that, either because of natural or government-imposed barriers, many goods are not traded. For nontradable goods there is obviously no reason for equalization of prices.

In fact, estimates suggest that fifty percent of most countries' output consists of nontradable goods. Considerable differences may arise when price inflation differs between the traded and non-traded goods sector. Froot and Rogoff (1995) devote careful attention to a hypothesis that deviations from PPP will arise due to the inclusion of non-traded goods in wholesale and consumer price indices. Their findings suggest that non-tradables are essential in explaining partial pass-through. Furthermore, Burstein *et al*. (2000) show that consumption goods contain distribution services around 47% in final price for agriculture sector and 42% in manufacturing. To the extent that differences in the efficiency of the distribution sector across countries remain constant over time, they would simply generate constant gaps in consumer price levels across countries. Similarly, to the extent that differences change over time, they would induce trends in relative prices.

CPI is the most widely used price index, probably because it is readily available and fairly comparable across countries. The information problem with trying to implement PPP using CPI is that governments do not construct indices for an international standardized basket of goods.[14] The calculations of PPP involve large amounts of nontradables and different baskets of goods in different countries. This is a serious problem, especially if developed and developing countries are compared. As pointed out

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[14] See the discussion in Summers and Heston (1991).

by Rogoff (1996) the US and German consumer price indices and producer price indices are conceptually quite similar. Although consumer price indices in different industrialized countries are conceptually quite similar, they are still constructed somewhat differently and the basket weights are not the same in any event. Sjaastad (1998) finds that measurement errors accounted for 75% of the variance in the real exchange rate.

Summers and Heston (1991) constructed a so-called ICP (international comparison program) data set to find a solution to these problems. They report estimates of absolute PPP for a long sample period and a number of countries, using a common basket of goods across countries. Since there are several problems in ICP, especially long time intervals and extensive use of extrapolation, official price indices still remain the basis commonly used in much empirical work. Recently, Xu (2003) compared the ability of alternative price indices to forecast the nominal exchange rate based on PPP. He finds that the choice of price indices greatly affects the quality of exchange rate forecasts. Among the three price indices, the CPI based forecasts are the worst.

The absolute version of the PPP hypothesis requires that the weights are equal in domestic and foreign price indices. Clearly, the greater the disparity between the relevant national price indices, the greater the apparent disparity from aggregate PPP, even when the law of one prices holds for individual goods. The problem is smaller for the price indices constructed using a geometric index. As pointed out by Sarno and Taylor (2002, p. 68), this is because the geometric price indices are homogeneous of degree one and the differences in weights across countries will matter less where price impulses affect all goods and services more or less homogeneously.

Imbs *et al.* (2002) show that the failure to account for cross-sectional heterogeneity in the dynamic properties of the typical price indices components substantially explains the slow mean reversion of PPP estimates based, for example, on CPI price indices. The speed of reversion to parity depends in all likelihood on good-specific characteristics, and thus is not homogeneous across sectors. If aggregate estimates are run under the premise

of a unique autoregressive coefficient, heterogeneity is pushed into the residuals. Imbs *et al.* (2002) show that failure to allow for these differences induces a positive bias in aggregate half-life estimates and corrected estimates are perfectly in line with the real exchange persistence derived in a model with plausible nominal rigidities. Furthermore, aggregation bias is most prevalent amongst traded goods where observed persistence is found to be largest. However, after correcting for small sample bias, Chen and Engel (2004) show that the half-life estimates indicate that hetoregeneity and aggregation bias do not help to solve the PPP puzzle.

4.2 ESTIMATION METHODS

One of the most important issues in the international economics literature concerns the role of the economic fundamentals in explaining exchange rate behavior. No existing model seems to be able to consistently explain both the tremendous short-term volatility and persistence in the real exchange rate. Lumsdaine and Papell (1997) have proposed the idea that dollar based exchange rates are best described as series of broken trends. However, allowing such segmented trends is approximately the same as excluding the 1980s and perhaps also the beginning of new millennium from the analysis. Instead, the more pressing question should be the sources of these stochastic trends.

According to Froot and Rogoff (1995, p. 1649) there are three different stages of empirical tests for PPP. The first stage includes a correlation-type test in which the null hypothesis is that PPP holds. The second stage involves unit tests test in which the null hypothesis is that deviations from PPP are completely permanent. The third stage consists of cointegration tests in which the null hypothesis is that deviations from any linear combination of prices and exchange rates are permanent.

In the context of cointegration analyses between variables, many empirical studies based on the Engle-Granger test do not find supportive evidence for PPP at the beginning of nineties. (e.g. Patel,1990; Kim, 1990). The low power of the Engle-Granger co-

integration test is often cited as a cause of the rejection of the PPP hypothesis. In a linear cointegration-based approach Johansen's maximum likelihood method allows testing in a multivariate framework, i.e. we do not need to place one variable on the left-hand side and use others as regressors as in Engel-Granger cointegration methodology. This is a very desirable feature since the test for cointegration should be invariant to the choice of the variable selected for normalization. Thus, we can consider the error structure of the data process allowing interactions in the determination of the relevant economic variables and also independently of the choice of the endogenous variable. Cheung and Lai (1994), for example, show that the long-run PPP between the US and UK, France, Germany, Switzerland, and Canada was supported based on the Johansen cointegration tests, but rejected based on the Engle-Granger tests.

4.3 UNIT ROOTS

The failure of PPP to hold continuously is well documented empirically (e.g. Froot and Rogoff, 1995 and MacDonald, 1995). While few economists believe that PPP holds at each point in time, most instinctively believe in some variant of purchasing power parity as an anchor for long-run exchange rates. (Rogoff, 1996, p.647). Attention focuses now on whether this variable is stationary, i.e. the validity of purchasing power parity hypothesis has been widely tested in empirical analysis of economic time series using unit root tests. Testing for unit roots is almost mandatory in the PPP literature.

The determination of the order of integratedness of a time series such as PPP is seldom unanimous. Theoretically we can classify variables exhibiting a high degree of time persistence as nonstationary I(1) variables and variables exhibiting a significant tendency to mean reversion as stationary I(0) variables, i.e. the variable is I(d) with d being 0, 1, or some greater integer.[15] In practice many variables, or a combination of variables, are borderline cases such that distinguishing between a strongly autoregressive I(0) or I(1) process is far from easy.

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[15] In general a time series can be fractionally integrated so that d need not be an integer.

There is a quite strong case to be made that stationarity or nonstationary is not a general property of an economic variable but a convenient statistical approximation to distinguish between the short-run, medium-run and long-run variation in the data. For instance, if the time perspective of the study is the macroeconomic behavior in the medium run, the real exchange rate probably exhibits considerable inertia, consistent with nonstationary rather than stationary behavior. From an econometric point of view the question remains in what sense a unit-root process can be given a structural interpretation.[16]

In general, the value of the exchange rate k period ahead can be written:

$$q_{t+k} = \overline{q}\sum_{i=0}^{k}\delta^i + \sum_{i=0}^{k}\delta^i\omega_{t-i} \; ,$$

where ϖ is a mean zero i.i.d. shock. The speed of adjustment to PPP depends on the value of parameter δ. Assuming that $\delta < 1$, the exchange rate is expected ultimately to converge on the constant, $\overline{q}$.

There are two popular unit root tests in the literature, namely the augmented Dickey-Fuller test (thereafter ADF) and the Philips-Perron test (thereafter PP). The basic property of these tests is that they assume nonstationarity as a null hypothesis. It has been shown that these tests lack power against meaningful alternatives, especially in small samples.[17] The test procedure called KPSS unit root test is often referred to as a more suitable unit root test than the ADF or PP test. It has a useful property that the hypothesis concerning stationarity holds under the null, and is rejected under the alternative in contrast to ADF and PP. It has, however, the same low power problem as the ADF and PP tests. Its usefulness for confirmatory analysis in conjunction with the ADF and PP tests could be a problematic case (with two tests that lack power).

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[16] See the discussion in Juselius (1999). Treating the real exchange rate as a nonstationary variable makes it possible to find out which other variables have exhibited similar stochastic trends.
[17] See Maddala and Kim (1998)

If the unit root model can characterize real exchange rate behavior, then PPP does not hold because there is no propensity to revert to the equilibrium level. A possible interpretation of the widespread failure to reject non-stationarity of real exchange rates is that the span of available data for a recent floating period may simply be too short to provide any reasonable degree of test power in the normal statistical tests for non-stationarity. Arguing that the post-Bretton Woods period may be far too short to reveal PPP reversion, many studies explore long historical data and find evidence of parity reversion in real exchange rates. Long-run data, for a century or more, which spans several exchange rate regimes, have been analyzed to improve the power of unit root tests. Lothian and Taylor (1996) discovered that the probability of rejecting a false null hypothesis was extremely low with 20 or even 50 years of annual data, but became acceptable over long spans. Cheung and Lai (1994) find evidence of mean reversion for WPI rates across several countries for the period 1900-1992. The long-horizon approach, however, is susceptible to a specific sample-selection bias (Froot and Rogoff, 1995).

4.4 PANEL UNIT ROOTS

The most popular recent method for circumventing the low power problem of unit root test is the use of the panel unit root method. The principal motivation behind panel data unit root tests is to increase the power of unit root tests by increasing the sample size. Applying such a method has typically allowed for the production of more evidence for real exchange rate mean reversion (e.g. Frankel and Rose 1996, MacDonald 1996, Lothian, 1997).[18] Panel unit root tests, however, are not free from potential drawbacks including their excessive sensitivity to country groupings and panel size.

Pappel and Theodoridis (2001) investigated the implications of the choice of numeraire currency on panel tests of PPP under the current regime of flexible exchange rates. They show that the conditions necessary for numeraire irrelevancy are not supported

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[18] Exceptions are, among others, O'Connell (1998) and Chortareas and Driver (2001).

empirically, and that the choice of numeraire currency can and does matter for PPP. The evidence of PPP is stronger for European than for non-European base currencies. Distance between the countries and volatility of the exchange rates are the most important determinants of the results.

The validity of panel unit root tests in the investigation of PPP depends on the hypothesis of interest. As pointed out by Maddala and Kim (1998) one may be interested in testing whether the hypothesis of the PPP holds for a certain bilateral exchange rate. In this case it is no use to be told that we reject the validity of the PPP even in the long run for this bilateral exchange rate but that if we throw in a large number of countries and use panel unit root tests, we do not reject the PPP hypothesis for this exchange rate. (e.g. MacDonald (1996). If the hypothesis of interest is, for example, an estimate of half-life of deviation from PPP, for this purpose the use of panel data is an appropriate procedure.

Since the conventional panel unit root tests assume no cross-sectional correlation, they cannot be directly applied to testing for reversion in exchange rates, since by construction that assumption is violated. Taylor and Sarno (1998) illustrate that joint nonstationarity of a group of real exchange rates may be rejected when only one of the series is mean-reverting. The analysis in Cheung and Lai (1998) uncovers significant heterogeneity in the behavior of real exchange rates across countries.

While studies utilizing panel procedures or long spans of data have generally been successful in rejecting the unit root hypothesis for real exchange rates, these studies also found that deviations from PPP are very persistent. Although unit root findings are heavily time span dependent, very slow mean reversion is a serious challenge for the real exchange rate literature. Thus, the key to resolving the possible failure of PPP for a recent floating period lies in understanding the forces that keep real exchange rates away from parity.

4.5 NONLINEAR UNIT ROOT TESTS

The degree of persistence in the real exchange rate can be used to infer the principal impulses driving exchange rate movements. This is crucial for many dynamic open-economy macro models, since the implications of those models are very sensitive to the presence or absence of persistent stochastic trends in real exchange rates (Lane, 2001).

Much of the controversy regarding the usefulness of the purchasing power parity doctrine is due to the fact that the doctrine does not specify the precise mechanism by which exchange rates are linked to prices. In standard linear cointegration methodology (Engle-Granger combined Dickey-Fuller test or Johansen procedure), the speed of adjustment to restore equilibrium is independent of the magnitude of disequilibrium. To find new insights into the persistency issue the recent literature explores possible nonlinearity in the speed of PPP reversion. The notation that real exchange rates could follow nonlinear processes dates back to Heckscher (1916), who suggested that deviations from the law of one price might be due to international transaction cost between spatially separated markets. Indeed, recent work indicates that while the random walk is a reasonably good approximation for short-run dynamics, real exchange rates show mean-reverting tendencies over the medium to long term.

The linear AR(1) specification of the standard Dickey-Fuller unit root model assume that reversion occurs monotonically, regardless of how far the process is from parity. In general, the augmented Dickey-Fuller equation is stated as follows.

$$\Delta\mu_t = k + \lambda\mu_{t-1} + \sum_{i=1}^{p-1}\gamma_i\Delta\mu_{t-i} \qquad\qquad (4.1)$$

where μ_t is a real exchange rate, k is a constant and λ, expected to vary between zero and minus one, is the convergence speed. A famous half-life of deviation from the parity level is determined as $\ln(0,5)/\ln(1+\lambda)$. The half-life measures mean reversion defined as

the number of time periods it takes for deviations to subside permanently below fifty per cent in response to a unit shock in the level of the series.

Consider the following non-linear specification of augmented Dickey-Fuller test.

$$\Delta\mu_t = k + \lambda\mu_{t-1} + \sum_{i=1}^{p-1}\gamma_i\Delta\mu_{t-i} + (k^* + \lambda^*\mu_{t-1} + \sum_{i=1}^{p-1}\gamma_i\Delta\mu_{t-i})G(\mu_{t-d}) + \varepsilon_t \qquad (4.2)$$

The larger the deviation from the parity level, the stronger the tendency to adjust to the parity level. 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. The model may also be viewed as a nonlinear error correction model in the form of a smooth transition autoregressive process.

The most common transition function is

$$G(\gamma, c, s_t) = 1 - \exp\{-\gamma(s_t - c)^2\} \qquad\qquad \gamma > 0. \qquad (4.3)$$

The model is called the exponential smooth transition regression model (ESTAR). s_t is the transition variable.[19] The stationarity of the switching variable is essential, because it is necessary that the process visits every regime infinitely often. If the switching variable is not stationary, the process has a certain probability to be absorbed into a single regime (Bec *et al.*, 2002, p. 3). The slope parameter γ indicates how rapid the transition from zero to unity is as a function of s_t. Finally, c is the location parameter, which determines where the transition occurs.

The transition function is symmetric about c and $G(\gamma, c, s_t) \rightarrow 1$ for $s_t \rightarrow \pm\infty$. This is a suitable assumption if, for example, we assume that the non-linearity is due to symmetric

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[19] The ESTAR model has been applied to real exchange rates by Michael *et al.* (1997), Taylor *et al.* (2001) and Baum *et al.* (2001).

and proportional transaction costs. The nonmonotonic second-order logistic function (LSTAR2) enables consideration of an asymmetric mean reversion toward parity level as well as of a sudden regime change.

$$G(\gamma, c, s_t) = (1 + \exp\{-\gamma(s_t - c_1)(s_t - c_2)\})^{-1} \tag{4.4}$$

Typically, the range of ESTAR type transition functions values indicates that convergence to long-run PPP is low, especially in the post-Bretton Wood era.[20] An analysis of the impulse response functions will allow the half-life shocks to the real exchange rate models to be gauged more precisely. According to the results of generalized impulse response functions in Baum *et al.* (2001) the speed of adjustment towards parity level is low and positive and negative shocks of the same magnitude appear to have different dynamic effects, thus suggesting sign asymmetry based on the sign of the shock.

Taylor *et al.* (2001) also reports impulse response functions corresponding to their estimated nonlinear real exchange rate models. The estimated half lives of four major bilateral real exchange rates illustrate the nonlinear nature of the response to shocks, with large shocks mean reverting much faster than smaller shocks. The dollar-mark, for example, displays quite fast mean reversion, ranging from a half life of under one year for the largest shocks to under three years for small shocks.

Rapach and Wohar (2003) find rather limited evidence of nonlinear behavior in US dollar real exchange rates. They re-examine the fitted Band-TAR models in Obstfeld and Taylor (1997) and the ESTAR models in Taylor *et al.* (2001) using post Bretton Wood data and find little difference in the conditional expectations functions for the fitted nonlinear models and linear counterparts. The results in Michael *et al.* (1997), which uses long span data, has been found to be more robust.

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[20] Michael et al. (1997) find out stronger convergence in the 1920s and in the two-century time period.

When autoregressive models are used, standard estimators, such as least square estimators, are significantly downward biased in finite samples, as noted by Cashin and McDermott (2003). They carefully study the importance of a median unbiased estimator in half-life estimates. The results in Cashin and McDermott (2003) show that while median unbiased estimators always increase the estimated half life of deviations from PPP in comparison to those derived from conventional methods, there is evidence of slow reversion of real exchange rates towards parity, which is consistent with PPP holding in the post-Bretton Woods period.

5. ESSAYS ON THE PPP PUZZLE

5.1 Long-run deviations from the purchasing power parity between the German mark and the U.S. dollar: Oil price-the missing link?

The aim of the first essay is to identify and investigate empirically the long-run determinants of real exchange rate fluctuations between Germany and the United States since the collapse of the Bretton Woods fixed exchange rate system. A number of studies, such as Engel (1999) and Canzoneri *et al*. (1999), have furnished fairly persuasive evidence that the deviations from purchasing power parity derive in large part from differences in relative prices, especially if the US dollar is included in the vector of time series. These findings have led some researchers to suggest that there might be yet an unidentified factor causing persistent shifts in the real exchange rate. We show that a positive oil price shock both appreciates the US dollar real exchange rate and also decreases the pass-through of changes in the exchange rate to consumer prices. A reduction in the pass-through is based on increased uncertainty related to the permanence of shock.

A typical explanation in the international financial literature is that the rising oil price leads to increased demand for dollars by foreign currency-area buyers. We build, however, our explanation on current and capital accounts. Whether the net effect is

favorable or unfavorable for the dollar depends on whether OPEC investments in dollars are greater or less than America's share of the industrial world current account.

It is difficult to rationalize a large share of the US assets relative to the German assets in the OPEC portfolio if agents are only interested in minimizing the risk for any given level of return. This is because the covariance of returns is high. In practice, it may be impossible to separate the economic and political considerations underlying investments, i.e. investments in dollar assets are not purely based on economic considerations but also on political issues. Thus, the central role of the U.S. in the Middle East might also affect investment decisions. Including the oil price in the observation vector, makes it possible obtain positive evidence for the traditional Harrod-Balassa-Samuelson effect not generally found between the German mark and the US dollar.

5.2 The U.S. dollar real exchange rate. A real options' approach

The major dollar appreciation of the 1980's caused a huge decline in the dollar price of traded goods sold in foreign countries relative to the dollar price of traded goods sold in the US. The high degree of correlation between the nominal exchange rate and the real exchange rate was almost complete, i.e. there was very little adjustment in the nominal prices of traded goods.

The aim of this paper is to discuss the determinants of the U.S. dollar real exchange rate fluctuation. We focus our analysis on the exchange rate effect on tradable prices. The disconnection between exchange rates and prices is rationalized using a real option theory following Dixit (1989a and b). Dixit (1989a and b) assumes that there are sunk costs of arbitrage and that the exchange rate follows a geometric Brownian motion process. The expected uncertainty is introduced using the historical variance of the exchange rate, i.e. it is assumed to be an almost constant variable over time. The inaction band (no entry or exit) around the base value is either stable or determined by the market share of foreign firms.

Although the values given for entry and exit trigger points in Dixit (1989b) are realistic especially in the median firm case, we argue that the dynamic structure of the pass-through process is at least partly problematic. We argue that the US dollar exchange rate time series process is better characterized by the mixed Brownian motion Poisson jump process than the traditional continuous time series process. Deviations from the parity level are too volatile and persistent to be explained solely by the changes in number of firms or small changes in expected uncertainty if uncertainty is generated by the continuous time series process.

We explicitly consider the effects of profit maximizing foreign firms' entry decisions on the domestic tradable prices through the supply changes after a large appreciation. The results of the theoretical analysis show that co-movements of prices and exchange rates do not depend only on demand and cost parameters, but also on the stochastic process followed by the exchange rate. Foreign suppliers do not completely adjust their supply after a large appreciation if there is a substantial likelihood of a large negative shock. Thus, the sensitivity to initial conditions is a crucial feature of the complex dynamics of arbitrage. We are able to explain large and persistent deviations from the parity level which are not constant in magnitude without assuming market imperfections or a systematic link between trade flows and the deviations of the real exchange rate from trend.

5.3 The Purchasing power parity puzzle: A sudden nonlinear perspective

The aim of this paper is to construct a simple nonlinear model for the U.S. dollar - German mark (euro) real exchange rate over the period 1975 to 2003. An exchange rate is an important relative price, one that potentially feeds back immediately into a large range of transactions. Exchange rates, however, are much more volatile relative to any model

we have of underlying fundamentals. The empirical evidence strongly suggests that the exchange rates of the major currencies are disconnected from fundamental economic variables most of the time. Unanticipated shocks in the fundamental variables explain only a small fraction of the unanticipated changes in the exchange rates. Typically over horizons of up to one year, news on output, inflation and interest rates explains less than 5% of the total variance of the exchange rate.

If the real exchange rate follows a random walk, then innovations to the real exchange persist and the time series may fluctuate without band. Recently, several authors (e.g. Michael, *et al.,* 1997; Taylor *et al.,* 2001) have argued that real exchange rates follow a nonlinear process. Intuitively, nonlinearities in the real exchange rate imply that frictions in international trade result in bands within which relative international prices can fluctuate without a strong tendency to adjust towards the parity level. Thus, goods-market arbitrage should prevent the exchange rate from fluctuating without boundaries.

PPP, however, is somewhat distinct from the pure law of one price concept applied to commodities. Our view is that PPP is examined more properly in the domain of monetary economics and macroeconomic theory of inflation. The weight placed on the exchange rate in a monetary policy rule is one of the main decisions facing the central bank in a modern open economy.

We assume that both central banks, the Fed and the Bundesbank (the ECB), respond to deviations in the expected inflation and output from their desired levels, i.e. both central banks follow a Taylor rule. However, the indirect effects of exchange rates on interest rates through standard domestic variables guarantee that the exchange rates do not depart very far from purchasing power parity. The weight put on the real exchange rate in the monetary policy rule is positively related to the pass-through from the exchange rate to import prices. Frictions in international trade imply that the weight put on the real exchange rate in the monetary policy rule may be lower close to PPP than far away from it.

As discussed in Kilian and Taylor (2003), when the exchange rate is far from latent equilibrium the consensus will be gradually built among fundamentalist that the exchange rate is misaligned. We will argue, however, that even though individual market agents may firmly believe the exchange rate to be misaligned, they hardly have the power or resources to buck the trend. Then an unanticipated monetary policy shock may be necessary to break a trend. Sarno and Taylor (2001) suggest that even sterilized interventions might be useful once the exchange rate has moved a long away from the fundamental equilibrium. Publicly announced interventions can now be seen as fulfilling a coordinated role in that they encourage individual market agents to enter the market at the same time.

It is reasonable to assume that the central bank is expected to extend only to monetary shocks that do not threaten its primary objective of low deviations of domestic target variables from their desired levels. Thus, publicly announced interventions signal future monetary policy intentions only if the exchange rate is relatively far away from parity level. If this distance from parity level is stable, market agents may recognize it. This implies a homogeneous response to large deviations from the PPP equilibrium.

In our empirical work, we do not primarily attempt to identify and trace through the effects of expected monetary shocks. Instead, we do aim to carefully elaborate the idea that the low weight put on the real exchange rate in the monetary policy rule in the neighborhood of PPP results in bands within which the deviations from PPP are possible. Based on the assumption of heterogeneous agents, smooth transition models are considered. Since the role of the central bank may be crucial, it is desirable to allow the adjustment towards long-run equilibrium to be discrete as well as smooth. We use a nonmonotonic second-order logistic smooth transition autoregressive model (LSTR2) to investigate nonlinear mean reversion. Thus, the model presented here departs from the conventional (exponential) nonlinear models due to different transition function.

We found that the adjustment is sudden and symmetric around the parity level. This may indicate that near the lower and upper boundaries the expectations on future monetary

reactions are increased, i.e. traders have homogeneous expectations near the boundaries. This is in all probability due to official announcements by the central bank(s). Furthermore, the band around the PPP equilibrium is relatively large. This implies that the bilateral dollar-mark exchange rate seems to fluctuate within a wide and flexible reference band.

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Long run deviations from the purchasing power parity between the German mark and the U.S. dollar.

Oil price-the missing link?

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Markus Lahtinen: **Long run deviations from the purchasing power parity between the German mark and the U.S. dollar. Oil price-the missing link?**.

The aim of this paper is to identify and investigate empirically the long-run determinants of real exchange rate fluctuations between Germany and the United States since the collapse of Bretton Woods system. This study uses the multivariate cointegration technique to find out steady state relations of the real exchange rate. Possible I(2) property of the data generating process is also examined using the latest statistical tools.

A number of studies have furnished fairly persuasive evidence that the deviations from the purchasing power parity (PPP) derive in large part from differences in relative traded goods prices especially if the U.S. dollar is included into the estimation vector. These findings have led some researchers to suggest that there may be an unidentified real or financial factor causing persistent shifts in the real exchange rate. In this study we have used a more explicit explanation for this exchange rate shock. Real oil price is shown to be an important factor for the real exchange rate movements between the German mark and the U.S. dollar. When the oil price shock is included into the estimation vector, we can also identify the classical Harrod-Balassa-Samuelson condition not generally found between the German mark and the U.S. dollar.

1. INTRODUCTION

The purchasing power parity (hereafter PPP) hypothesis states that national price levels expressed in a common currency should be equal. The empirical evidence is, however, rather contradictory. Recent evidence suggests that the failures of the law of one price are not only significant, but they also play a dominant role in the behavior of real exchange rate. The aim of this paper is to identify and investigate empirically the long-run determinants of real exchange rate fluctuations between the United States and Germany since the collapse of the Bretton Woods system of fixed exchange rates in the early 1970's. The real exchange rate is measured using CPI deflators.

Although the number of previous attempts at modeling real exchange rates for recent floating period has been enormous, such a work has not always proved particularly fruitful.[1] Bilateral exchange rate models fails typically to establish a significant long-run link between the real exchange rate and fundamentals. However, one key message to come from existing research on the modeling exchange rates is that the econometric method used can have a crucial bearing on the findings of significant and sensible long-run relationships. This paper uses the multivariate cointegration technique developed by Johansen (1988, 1991) to find steady state relations of real exchange rates. The Johansen's maximum likelihood approach has been shown to be a more efficient method to analyze cointegration structure relative to the traditional single equation method of Engle and Granger.[2]

Allowing modeling in a multivariate framework is one of the most crucial advantages of the Johansen maximum likelihood approach. Furthermore, the possibility to impose restrictions on the cointegration vectors is important, especially if the model contains more than one cointegration vector. In this paper we follow a research tradition inspired by Johansen and Juselius (1992). All the variables are included in levels and possible cointegrated relations are examined by an analysis of the likelihood function.

One important theme in this paper concerns the importance of distinguishing between statistical and theoretical measures of a unit root. There are many arguments in favor of considering a unit root (a stochastic trend) as a convenient econometric approximation rather than as a deep structural parameter. For example, in many empirical studies inflation is assumed to follow a stationary

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[1] As a survey, see Rogoff (1996).
[2] See the discussion in Maddala and Kim (1998).

process. Although this is theoretically only acceptable assumption in the very long run, in the medium run it may not be a suitable statistical formulation, i.e. for a period such as the recent float after Bretton Woods. For a period such as the recent float prices are found to be even $I(2)$, implying that inflation rates are $I(1)$. Possible $I(2)$ property of the data generating process is examined using the latest statistical tools developed by Johansen (1995), Paruolo (1996) and Rahbek *et al.* (1998). This analysis gives us a possibility to examine more closely the relationship between the prices and the nominal exchange rate.

Another important issue concerns the persistence in real exchange rates. The main theoretical explanation for the phenomenon that real exchange rates do not have unit roots in the long run is that if the predominant force upsetting the PPP relationship is nominal then this will have only a transitory effect on deviations from PPP (as in the celebrated Dornbusch over-shooting model). If money is neutral in the long run, the real exchange rate should be a stationary variable. A number of studies have demonstrated that for the recent floating experience real exchange rates are $I(1)$ process. This evidence suggests that there are also exogenous real shocks which may affect the deviations from the PPP. If the sources of disturbances are real in nature, we would argue this will have a more permanent effect on the real exchange rate.

Sectoral productivity differentials across countries have long been suggested as major determinants of real exchange rate movements in the long run. This effect, known as the Harrod-Balassa-Samuelson (hereafter HBS) effect, is a tendency for countries with higher productivity in tradables compared with non-tradables to have higher price levels.[3] To be more accurate, the HBS hypothesis divides real exchange rate movements into two components. The first component of the hypothesis is the assumption that the relative price of non-tradables is proportional to the ratio of the marginal product of labour in the tradable sector to that in nontradable sector and the second component is the assumption of PPP for traded goods. These two components combine to produce a simple model of real exchange rate movements.

A number of studies, such as Asea and Mendoza (1994) and De Gregorio and Wolf (1994), have furnished fairly persuasive evidence that, at least for industrial countries, deviations from PPP derive in large part from differences in relative traded prices across countries. Engel's (1999) results concerning the U.S. dollar are similar and they suggest that consumer prices for tradables goods behave very much in the same way as non-tradables consumer prices. These results are quite

puzzling in a view of HBS theory which holds that the relative tradables prices show little long-term variation across countries compared with the variation in relative non-tradables prices. However, as discussed in MacDonald (2000), this does not necessarily imply that the HBS effect is in itself unimportant or insignificant, but the above evidence just suggests that the dominant component of the real exchange rate behavior is the nominal exchange rate even in the long run.

We find that oil price appreciates the US dollar exchange rate relative to the German mark.[4] The importance of oil price for the US exchange rate movements has been argued by Krugman (1983). This model builds on long-run changes on the balance of payment due to the change in oil price using a multi country framework. Higher oil price will transfer wealth from the oil importers to the oil exporters. While current accounts are thus worsened, there is an improvement in capital accounts as oil exporters invest their trade surplus in foreign currencies. Whether the net effect is favorable or unfavorable for the US dollar depends on whether investments in dollars are more or less than America's share of the industrial world's current account deficit. Trade flows, however, are more important in the long-run. The net improvement in the U.S. balance of trade would then also require a real appreciation of the US dollar.

Froot and Klemperer (1989) show that temporary changes in the nominal exchange rate may have a relatively small effect on price differentials. Oil price shock may be an important source of large and temporary changes in the nominal exchange rate. When the oil price shock is included into the observation vector, we can identify a classical Harrod-Balassa-Samuelson condition not generally found in the literature between the German mark and the U.S. dollar.

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[3] See Balassa (1964) and Samuelson (1964).

[4] The relationship between oil prices and real exchange rates has also been examined, for example, in Johansen and Juselius (1992), Rogoff (1992), MacDonald (1997), and Amano and Van Norden (1998). Johansen and Juselius introduce oil price as an exogenous but significant variable for the PPP relation of the United Kingdom. Rogoff (1992) finds oil price a significant variable for the Japanese yen-the U.S. dollar real exchange rate. However, this finding depends on the chosen time period. MacDonald (1997) finds a weak support for the importance of oil price using a multivariate cointegartion method for the real effective U.S. dollar exchange rate, the German mark, and the Japanese Yen. Amano and Van Norden (1998) use the U.S. dollar real effective exchange rate and find oil price significant variable even in a bivariate system.

2. TRADABLE AND NON-TRADABLE GOODS

The real exchange rate is a measure of one country's overall price level relative to another country's. The real exchange rate, defined in terms of the general or overall price level, such as the CPI, is given by

$$q_t = p_t - p_t^* - s_t \qquad\qquad (2.1)$$

where q_t denotes the logarithm of the real exchange rate, p_t denotes the log of the domestic price level, p_t^* the log of the foreign price level and s_t the log of the nominal exchange rate defined as the home currency price of a unit of foreign currency. In this context, therefore, a rise in q_t denotes an appreciation of the real exchange rate. To measure the price level, we decompose it into the traded and non-traded components and use a geometric average of these prices in both country

$$p_t = (1-\alpha)p_t^T + \alpha p_t^N, \qquad \alpha < 1 \qquad\qquad (2.2)$$

where p_t denotes the logarithm of the price index, p_t^T is the log of the traded goods price index, p_t^N is the log of the non-traded goods price index and α is the share that nontraded goods take in the price index. Letting an asterisk represent the foreign country, one can also write

$$p_t^* = (1-\beta)p_t^{T*} + \beta p_t^{N*} \qquad \beta < 1 \qquad\qquad (2.3)$$

where β is nontraded good's share in the foreign price index.

2.1. NON-TRADABLE PRICES

In order to define the Harrod-Balassa-Samuelson condition (HBS) we have to assume perfect international competition of goods and capital markets to ensure that the prices of tradables and interest rate are pinned down.[5] The former then determines uniquely the wage rate of internationally immobile assumed labor by equalization of marginal product and the given world price. This, with given intersectional factor mobility (labor and capital), means that relative prices are set exclusively by the level of productivity in the two sectors, i.e. the productivity in the tradable sector then determines also the price of non-tradables.

Since labor and capital factors are free to move between sectors costlessly, only supply side factors matter.[6] In the HBS model demand side factors will affect the real exchange rate if, for example, the assumption of perfect competition, PPP for traded goods, or perfect capital mobility is relaxed. The conclusion is, ceteris paribus, that economies with a higher level of productivity in tradables will thus be characterized by higher wages and also by higher price of non-tradables if productivity in a non-tradable sector does not increase in same extent, i.e. economy will face a more appreciated real exchange rate.

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[5] This is completely true only in a small open economy.
[6] See also the discussion in Dornbusch (1989). Necessary conditions for HBS are obviously more acceptable in the long-run.

2.2. TRADABLE PRICES

As discussed in Obstfeld and Taylor (1997), even the studies most favorable to long-run PPP suggest an extremely slow decay rate for international price differentials. Estimated half-lives for PPP deviations for most countries and time periods are found to be of the order of four to fives years.[7] These estimates appear to imply more sluggishness than one can attribute entirely to nominal rigidities alone.

Engel's (1999) results suggest that consumers' prices for tradables goods behave very much in the same way as non-tradables consumer prices. These results are quite puzzling in view of the Harrod-Balassa-Samuelson theory which holds that the relative tradables prices show little long-term variation across countries compared with the variation in relative non-tradables prices. It is certainly true that the non-tradable component is important in determination of tradables prices. However, Engel (1999) results again make clear that more than just this must be going on, since prices for relatively tradable goods do not seem to respond any faster to exchange rate movements than do the prices of non-traded goods.[8]

The latest generation of studies on prices and exchange rates has focused more sharply on the issue of markup adjustment as a possible explanation for very slow response of tradable goods prices to exchange movements. Krugman (1987) labeled the phenomenon of exchange rate induced price discrimination in international markets "pricing-to-market", hereafter PTM. According to the PTM approach international markets for manufacturing goods are sufficiently segmented that producers can, at least over some horizon, tailor the prices they charge to the specific local demand conditions prevailing in different national markets. Although there is a large body of literature suggesting that PTM is indeed important to PPP deviations its implications for large and persistent PPP deviations are not clear.[9]

Trade frictions, such as transportation costs, allow even tradables prices to differ within some range without inducing profitable arbitrage. However, this no-arbitrage range is typically assumed to be narrow.[10] Something important must be also happening between the consumer level, where the

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[7] See Frankel and Rose (1996).
[8] Engel (1999) uses aggregate price indices, but there are also similar findings for highly disaggregated data. See Engel and Rogers (1996).
[9] See Goldberg and Knetter (1997) as a survey.
[10] Estimates of average transport costs across all tradable goods range between 5 and 10 percent.

medium-term effect of exchange rates on prices is virtually zero for many goods, and the wholesale level, where price effects tend to be much less than proportional but also significantly greater than zero.[11] According to Obstfeld and Rogoff (2000), apparent stickiness in terms of domestic currency of the import prices consumers face could result from the pricing practices of domestic importers and distributors, who purchase goods denominated in foreign currency but set retail prices in domestic currency. In that case, importing firms face international prices but the decisions of the ultimate consumers face the retail prices, i.e. not directly international terms of trade prices.

Froot and Klemperer (1989) show that a model with consumer switching costs will lead exporters to respond differently to temporary and permanent changes in exchange rate. They examine the effects of temporary appreciation of the dollar focusing on dynamic demand side effects in an oligopolistic market. In their model temporary appreciation increases the value of current, relative to future, dollar profits expressed in foreign currency. When the dollar is temporary high, foreign firms will find investments in market share less attractive, and will prefer instead to let their current profits to increase.

3. ECONOMETRICS OF REAL EXCHANGE RATE

A number of studies have demonstrated that for the recent floating experience real exchange rates are I(1) processes.[12] There are real fundamentals, such as productivity differentials, which may be responsible for introducing stochastic trends into real exchange rates. This interpretation has received some empirical support from researchers who have explicitly modeled the real determinants of real exchange rates.[13] However, statistical specifications of these studies are sometimes at least dubious. As discussed in Chinn and Johnston (1996), the pitfall of these models is that although they do tend to capture significant Harrod-Balassa-Samuelson links, they often rely on difference specification for bilateral and multilateral rates. Such tests are likely misspecified because the Harrod-Balassa-Samuelson hypothesis is about the relationship between the level of productivity and the level of the real exchange rate.[14] That is to say, if the series are I(1) then the Harrod-Balassa-Samuelson hypothesis implies that series must be cointegrated and therefore the

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[11] See the results in McCarthy (1999).

[12] Generally, the most convincing support for stationary real exchange rates is based on panel unit root tests. See, among others, Papell and Theodaris (1998), Koedijk et al. (1998), and Oh (1996).

[13] See Cheung and Lai (1993), MacDonald (1993), Amano and Van Norden (1995) and Canzoneri *et al.* (1999).

[14] See, for example, Hsieh (1982) or DeGregorio and Wolf (1994).

regression, which solely relies on differences will be mispecified from a statistical perspective. It is also important to observe that estimation in the first differences is consistent with the view that there is no meaningful concept of reversion to the productivity-determined equilibrium exchange rate.

A more common shortcoming of real exchange rate models in a multivariable case is that prices are usually only implicit variables and statistical properties of the data are not fully recognized. This leads again to the use of differenced time series to account for the (1) property of data, but then all long-run information in the levels of prices and exchange rates have been removed by differencing.[15] In this paper we follow a research tradition inspered by Johansen and Juselius (1992). All the variables are included in levels. Prices are included explicitly not only implicitly in a real exchange rate term. Possible I(2) trend generated by prices will be also analyzed.

The multivariate cointegration technique developed by Johansen (1988, 1991) is used in this study to find steady state relations of real exchange rates. Johansen's full-system maximum likelihood estimation technique for cointegration testing is based on VAR representation of time series. We consider both the cases of I(1) and I(2) in data generating process. By allowing for a set of conditioning variables, D_t, to control for institutional factors, and assuming multivariate normality, the vector autoregressive model is obtained as a tentative statistical model for the data generating process.

Let X_t be a $p \times 1$ vector of I(1) variables in the system. Since the basic idea of Johansen's method is to distinguish between stationarity by linear combinations and by differencing, we write the model in the following error correction form

$$\Delta X_t = \mu_0 + \mu_1 t + \Gamma_1 \Delta X_{t-1} + \Pi X_{t-1} + \Phi D_t + \varepsilon_t, \tag{3.1}$$

where ε_t is distributed $Niid(0,\Sigma)$ and the parameters $\Theta = \{\Gamma_1, \Pi, \mu_0, \mu_1, \Sigma\}$ are unrestricted. The parameter Γ_1 defines the short-run adjustment to the changes of the process. The matrix Π is estimated by Johansen maximum likelihood procedure subject to the hypothesis that Π has reduced rank, i.e. $\Pi = \alpha\beta'$, where α and β are p x r matrices, p > r. If r < p then under certain conditions the

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[15]Juselius (1999) and Juselius and MacDonald (2000) have, among others, discussed in this theme.

process ΔX_t is stationary, X_t is nonstationary, but also $\beta' X_t$ is stationary.[16] Thus we can interpret the $\beta' X_t$ as the stationary relations among nonstationary variables, i.e. cointegration relationships.[17]

Johansen (1991) showed that in addition to the restriction:

$$\Pi = \alpha \beta', \qquad\qquad (3.2)$$

where α defines the short run adjustment to the steady state relations, the following restriction also has to be satisfied

$$\alpha'_{\perp} (-I + \Gamma) \beta_{\perp} = \xi \eta' \qquad\qquad (3.3)$$

where ξ and η are (p - r) x (p - r) matrices, $\alpha_{\perp}$ and $\beta_{\perp}$ are p x (p-r) matrices orthogonal to α and β respectively. The parameterization of the restrictions facilitates the investigation of, on the one hand, the r linearly independent stationary relations between the levels of the variables, and, on the other hand, the p-r- linearly independent nonstationary relations.

If the second restriction for the I(1) model is violated, the process x_t is integrated of the second order or higher. When the process is I(2), it is useful to rewrite model in second order differences

$$\Delta^2 X_t = \mu_0 + \mu_1 t + \Gamma_1 \Delta^2 X_{t-1} + \Gamma \Delta X_{t-1} + \Pi X_{t-1} + \Phi D_t + \varepsilon_t . \qquad\qquad (3.4)$$

The hypothesis that X_t is I(2) is formulated as two reduced rank restrictions in Johansen (1991):

$$\Pi = \alpha \beta' \qquad \text{and} \qquad \alpha'_{\perp} \Gamma \beta_{\perp} = \xi \eta', \qquad\qquad (3.5)$$

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[16] Details concerning necessary conditions and also further analysis, see Johansen (1988).
[17] The real importance of the model formulation 3.2. is that it allows the precise formulation of a number of interesting economic hypothesis in such way that they can be tested.

where ξ and η are (p - r) $\times s_1$ matrices (s_1 is the number of I(1) trends), and the necessary and sufficient conditions for X_t to be I(2) are that rank (Π)=r < p, rank$\left(\alpha'_\perp \Gamma \beta_\perp\right) = s_1 <$ (p-r), and that a further rank condition holds which prevents the variables from being integrated of higher orders.[18] The linear trend coefficient μ_1 should restricted to $sp(\alpha)$, i.e. $\alpha'_\perp \mu_1 = 0$ as suggested by Rahbek *et al.* (1998) in order to avoid quadratic trends.

The space spanned vector X_t can be decomposed into r stationary directions, β, and p-r nonstationary directions, $\beta_\perp$, and the latter into directions, $\left(\beta_{\perp 1}, \beta_{\perp 2}\right)$, where $\beta_{\perp 1} = \beta_\perp \eta$ is of dimension p x s_1 and $\beta_{\perp 2} = \beta_\perp (\beta'_\perp \beta_\perp)^{-1} \eta_\perp$ is of dimension p x s_2 and $s_1 + s_2 = p - r$. It appears that both β and $\beta_\perp$ define nonstationary directions of the process, but that $\beta' X_t$ can be made stationary by a suitable combination of the differenced I(2) variables, whereas $\beta'_{\perp 1} X_t$ and $\beta'_{\perp 2} X_t$ can only be made stationary by differencing. Hence, even in I(2) model, the interpretation of the reduced rank of matrix Π is that there are r relations that can become stationary by cointegration (either directly or polynomially) and p – r noncointegrating relations can only become stationary by differencing.[19]

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[18] More detail discussion, see, Johansen (1992).
[19] Further discussion, see Juselius (1998).

4. EMPIRICAL MODELLING OF THE HARROD-BALASSA-SAMUELSON HYPOTHESIS

Based on the discussion presented in the section 2, we should find long-run cointegration relations using only the price and productivity variables, if PPP for traded goods holds. The idea is to begin our analysis with the Harrod-Balassa-Samuelson model and extend it to the full model as a second step of the analysis. Obviously, the initial information set should not to be too small to invalidate the identification of relevant cointegration relations. However, we are now testing a specific theoretical relation (the first component of Harrod -Balassa-Samuelson hypothesis) between chosen variables which makes even the result of "no cointegration" interesting. Later, the gradual expansion of the information set facilitates an analysis of the popular *ceteris paribus* assumption and its importance for empirical analysis of the basic Harrod-Balassa-Samuelson variables. Thus, the first information set analyzed gives rise to the following 5 variable model

$$X_t^{'} = (s_t, p_t, p_t^*, \mathrm{Pr}\,o_t, \mathrm{Pr}\,o_t^*)$$

where s_t is the nominal exchange rate, p_t is the German consumer price index, p_t^* the U.S. consumer price index, $\mathrm{Pr}\,o_t$ is the German productivity index, $\mathrm{Pr}\,o_t^*$ the U.S. productivity index. The data set used consists of quarterly time series observations from 1975:2 to 1998:1 for Germany and the United States. The main rationale for constructing the sample period to begin in 1975 is to abstract from any transition dynamics associated with the breakdown of Bretton Woods. All the variables are in logarithmic forms.[20]

As seen in the figures 1-10 in Appendix 1, the observations are strongly time dependent, pointing the need for models based on the adjustment to steady states. Therefore a probability formulation of the whole data set is needed. An unrestricted model was estimated for $X_t^{'}$. It is also important to find necessary dummy variables because, as Juselius (1994) has pointed out, adding more lags to the model is not a proper remedy for residual autocorrelation, if residual misspecification arises as a consequence of omitting important variables in a dummy vector. This may lead to heavily overparameterized models. The chosen VAR model needed seasonal dummies and the following two other dummy variables:

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[20] More detail description of the data is given in Appendix 1.

$$D_t^{'} = D91, D93,$$

where D91 is a dummy measuring (+1) in 1991:2 and 1991:3 and D93 (+1) in 1993:1. Thus, we have used a dummy variable especially designed for German reunification (D91) and a dummy variable measuring the effects of events related to the years of the European Exchange Rate Mechanism (ERM) crises in Germany (D93).[21] Both dummies are supported by the data set with the t-values 3,5 for D91 and 4,08 for D93.

The number of lags in VAR was increased until the residuals were Gaussian. Finally, the VAR was specified by installing two lags. Consequently, VAR(2) seems to provide a reasonable good approximation of the data generating process. Since all empirical models are inherently approximations of the actual data generating process, the question is whether our model is a satisfactory close approximation. To investigate this issue we test the stochastic specification regarding residual correlation, heteroscedastisity and normality. Test statistics are reported in Table 4.1 below. A significant test statistic is given in bold face.

Table 4.1. Misspecification tests.

Residual autocorr.	LM(1)	CHISQ(25) =	26,27	p-val =	0,39
	LM(4)	CHISQ(25) =	18,66	p-val =	0,81
Normality	LM	CHISQ(10) =	8,61	p-val =	0,57
Univariate tests	Ds	Dp	Dp*	Dpro	Dpro*
ARCH(2)	0,01	0,47	2,20	1,00	1,80
Normality	1,47	0,77	1,65	1,08	3,57
Skewness	-0,22	0,07	-0,31	0,18	-0,33
Ex. Kurtosis	0,22	0,12	-0,11	-0,45	-0,48
R-squared	0,23	0,71	0,81	0,43	0,47

Neither of multivariate tests are significant. This is an advantage for the model because the VAR model is based on the assumption of multivariate normal disturbances, i.e. residuals should behave approximately as a multivariate normal process. Univariate tests test for normality of the individual residuals can be rejected as a result of skewness (third moment) or excess kurtosis (fourth moment). Since the properties of the cointegration estimators are more sensitive to deviations from normality

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[21] Also the step dummy, 0 before reunification and 1 after, was examined but it did not work a satisfactory way.

due to skewness than to excess kurtosis we reported the third and forth moment around the mean.[22] Because there does not seem to be a serious skewness problem, we conclude that the normality conditions are satisfactory.

4.1 EMPIRICAL ANALYSIS OF THE I(2) MODEL

Selecting proper critical values for testing a cointegration rank depends on the nature of the deterministic components and the order of integration of the data. In the following analysis all test statistics have been calculated under the assumption that data contains linear but not quadratic trends. *A priori*, the differences between components in terms of persistence are due to the order of their stochastic trends rather than differences in the deterministic part.[23] The linear trend may, of course, have zero coefficients in certain directions. However, whether a trend is present does not affect the asymptotic properties of tests and estimators in chosen model.[24] In addition, because our data vector X_t might be of second order instead of first order nonstationary, the asymptotic distributions based on I(1) assumption might be violated. For example, asymptotic distributions of conventional tests used to find the correct inference on the number of cointegrating vectors, such as Trace and Max test, might be misleading if the order of integration is two.[25]

In this section we will first discuss the choice of rank based on the additional information given by the p x k =10 roots of companion matrix.[26] The number of unit roots in the characteristic polynomial is $s_1 + 2s_2$, where s_1 and s_2 are the number of I(1) and I(2) components respectively. The intuition is that the additional s_2 unit root belong to ΔX_t, hence, to the Γ matrix in (3.4). Therefore, the roots of the characteristic polynomial contain information on the unit roots associated with both Γ and Π, whereas the standard I(1) trace test only contains information on unit roots in the Π matrix. Additionally, if the choice of r incorrectly includes a nonstationary relation among cointegrating relations, then at least one of the roots of the characteristic polynomial of the model is a unit root or a near unit root. If there are no I(2) components, the number of unit roots should be $p - r$ and that is $p - r + 2s_2$ in the I(2) model.

[22] See the discussion in Gonzalo (1994).
[23] Doornik *et al.* (1998) found that even if the DGP did not include the trend its adoption into the cointegration space would only have a low cost.
[24] A property of asymptotic similarity, see Rahbek *et al.* (1998)
[25] Jörgensen (1998) demonstrates the low power of the trace tests in I(2) or near I(2) models.
[26] The discussion about characteristic roots and companion matrix see, for example, Kongsted (1998).

Table 4.2. The number of non-stationary trends.

	Five largest roots of the process				
Unrestricted model	0,97	0,97	0,95	0,89	0,48
r = 3	1,00	1,00	0,92	0,92	0,71
r = 2	1,00	1,00	1,00	0,95	0,73
r=1	1,00	1,00	1,00	1,00	0,95

The results reported in Table 4.2 show that the first root in the unrestricted model is the complex pair of roots with modulus 0,97 located almost on the unit circle followed by the roots with the modulus 0,95 and 0,89. The fifth largest root of unrestricted model, 0,48 is substantially smaller than the first four roots. Thus, this seems to indicate at most four roots in the data set. Imposing two unit roots into system, i.e. assuming r = 3, leaves two large unrestricted roots (0,92 and 0,92) in the model. However, imposing three or even four unit roots again leaves a large unrestricted root in the model. As Juselius (1998) has pointed out, a unit root in the characteristic polynomial that belongs to an I(2) trend cannot be removed by lowering r. Thus, our finding is a strong evidence for at least one stochastic I(2) trend. The results discussed here are consistent with one of the two following alternatives: (r = 3, $s_2 = 2$, $s_1 = 0$) or (r = 2, $s_2 = 1$, $s_1 = 2$).

The order of integration and cointegration can be formally tested in the I(2) model using the likelihood procedure. Johansen (1995) derived a LR test for the determination of s_1 conditional on chosen r. Paruolo (1996) extended the test procedure to the joint determination of (r, s_1) and Rahbek *et al.* (1998) derive the nonstandard asymptotic distributions for trend stationarity in the I(2) model. Two hypotheses given above were tested using this likelihood ratio test procedure. The test statistics reported in Table 4.3 are based on the VAR model with a trend in the cointegration space and, therefore, based on the tables in Rahbek *et al.* (1998). It is also defined that $\alpha'_\perp \mu = 0$, i.e. quadratic trends are not allowed in the model. The 95% quantiles are given in the lower part of Table 4.3. Note that the tabulated values are generated for a model without dummies and without small sample corrections. Therefore, the size of the tests is not likely to be accurate and the results should only be considered as indicative. In the following table a significant test statistic is given in bold face.

Table 4.3. Formal Test of I(1) and I(2) Cointegration Ranks.

p-r	r						Q(r)
5,00	0,00	387	290	220	173	140	131
4,00	1,00		241	152	105	**69**	67
3,00	2,00			131	**57**	**29**	27
2,00	3,00				86	**20**	14
1,00	4,00					35	3
	p-r-s$_1$	5	4	3	2	1	
p-r	r						Q(r)
5,00	0,00	198	168	142	120	101	84
4,00	1,00		137	113	92	75	63
3,00	2,00			87	68	53	42
2,00	3,00				48	34	25
1,00	4,00					20	12
	p-r-s$_1$	5	4	3	2	1	

The conventional test procedure starts with the most restricted hypothesis (r = 0, s_2 =5) in the upper left, and testing successively less and less restricted hypotheses according to Pantula (1989) principle until the first acceptance. It appears that the first acceptable structure to be (r = 2, s_2 = 2, s_1 = 1) indicating that at most two I(2) trends are supported by the data. The first acceptable structure of interest seems to be (r = 2, s_2 = 1, s_1 = 2). The second structure of interest (r = 3, s_2 = 2) is not supported by the data. Thus, we conclude that r = 2.

Inference in the I(2) model is based on asymptotic theory and there is not complete knowledge of infinite samples properties of cointegrating relations. Thus, the transformation to better known I(1) model is needed. A natural hypothesis which follows from the I(2) property of the prices is that the price differential is a first order nonstationary process, i.e. in the I(2) field, s_2 = 1 implies in this case that **p** and **p*** must contain the same I(2) trend and be CI(2,1) with the cointegration vector (1,1). However, this requirement needed for the transformation is rejected based on a test statistic of 25,4 distributed as $\chi^2(3)$. This finding is conformed by the results of I(2) test which indicates one I(2) trend even if a price differential transformation has been made (not reported).

If the long-run stochastic I(2) trend in prices is not the same for Germany and the U.S., resulting in a long-run I(2) trend in the price differential, then we would expect the nominal exchange rate to exhibit a similar long-run stochastic trend. The possible finding of I(2) nature of the nominal

German mark/US dollar exchange rate for this time period is not supported in a literature.[27] If the model is, however, estimated by assuming a common stochastic trend in both variables, there is no evidence on I(2) in the data. This is a very promising result and we will analyze this relation more closely.

In order to use the real exchange rate in the transformation vector we should first investigate a possible long-run homogeneity between variables included in the transformation. A necessary condition for the homogeneity is CI(2,1) between variables which presupposes the nominal exchange rate to be I(2). The hypothesis of long-run homogeneity can be tested as restrictions on β as well as its orthogonal complements $(\beta, \beta_{\perp 1}, \beta_{\perp 2})$ as described in section 3. The estimates $\beta_{\perp 1}$ define the CI(2,1) relations and $\beta_{\perp 2}$ define the variables which are affected by the I(2) trend.[28] The hypothesis of long-run homogeneity between chosen variables ($p_t - p_t^* - s_t$) can be formulated as:

$$\beta_i^{'} = [a_i, -a_i, -a_i, *, *], \quad i=1,\ldots,r \tag{4.1}$$

$$\beta_{\perp 1} = (b, -b, -b, *, *) \tag{4.2}$$

$$\beta_{\perp 2} = (c, c, c, 0, 0). \tag{4.3}$$

Because the real exchange transformation seems to eliminate the I(2) trend in the data we should see long-run price homogeneity assumption to hold when $(\beta, \beta_{\perp 1}, \beta_{\perp 2})$ directions are analyzed. These results depend heavily on the assumptions of the number of stationary and nonstationary relations. The only hypothesis which at least partly satisfies a long-run homogeneity assumptions is r $=2$ and $s_2 = 2$. However, the real exchange transformation itself may contain an I(1) trend. Assuming one I(1) trend and two I(2) trends is not in line with the number of roots in a companion matrix. There is now one extra unit root which we cannot find in a companion matrix. Because we do not completely understand finite sample properties of the I(2) model, especially when cointegration is a borderline case, the I(1) transformation is prioritized. Thus, the estimates reported in Table 4.4, are based on the assumptions r $=2$ $s_1 = 1$ and $s_2 = 2$, though admitting that the econometric evidence of the fifth unit root was not empirically robust.

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[27] See also the Figure 2 in Appendix 2. However, Juselius and MacDonald (1999) have made a borderline conclusion concerning the I(2) property of the nominal mark/dollar exchange rate during the recent float.
[28] See the discussion in Juselius and Toro (1999).

It is possible to test whether the long-run homogeneity assumption can be imposed in all cointegration relations (Hypothesis 4.1). The likelihood ratio test statistic 28,91 is asymptotically distributed as $\chi^2(4)$. Thus, and not surprisingly, the hypothesis concerning the overall long-run homogeneity between nominal exchange rate and price differential in the cointegration space is clearly rejected. The second hypothesis tests whether the real exchange transformation will lead to an I(1) model. This hypothesis is accepted at five percent significance level with the likelihood ratio test statistic 4,91 $\chi^2(3)$. In Table 4.4 the estimates for $\beta_{\perp 1}$ and $\beta_{\perp 2}$ are given.

Table 4.4. Estimates of $\beta_{\perp 1}$ and $\beta_{\perp 2}$ directions.

	s	p	p*	Pro	Pro*
$\beta_{\perp 1}$	3.04	-10.05	5.17	-0.27	-0.92
$\beta_{\perp 2}$	-2.74	-4.36	-2.94	-0.8	-0.76
$\beta_{\perp 2}$	-1.13	2.68	3.71	-1.22	-0.94

The result concerning $\beta_{\perp 1}$, defining the variables in CI(2,1) relation, are quite satisfactory. These results seem to suggest the relation between the price differential and the nominal exchange rate. Thus, it is likely that $p_t - p_t^* - s_t$ is CI(2,1), but not with the unitary coefficient in all cointegration relations.

If the cointegration property C(2,1) is accepted then all these variables should be I(2) variables. This assumption is partly supported by $\beta_{\perp 2}$ vectors which determine the weights with which the I(2) trend component influences the variables of the system. The time series behavior of the nominal exchange rate is difficult to interpret. These contradictory findings are probably attributable to a weak relationship between the nominal exchange rate and the price differential.[29] We conclude that there appears to be a weak cointegration relation between the price differential and the nominal exchange rate, although it does not strictly fulfill restrictions based on standard economic theory.

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[29] Note that we operate with linear models. Recent evidence in an international finance literature implies that the relationship between nominal exchanges rate and price differentials might be non-linear. Nonlinear models predict that nominal exchange rates and price differentials are only weakly related in a neighborhood of parity level. See Taylor, et al. (2001).

4.2. EMPIRICAL ANALYSIS OF THE TRANSFORMED I(1) MODEL

The empirical analysis of the first I(1) data set will be based on the real exchange rate transformed vector:

$$(ppp_t, \Pr o_t, \Pr o_t^*, \Delta p_t, \Delta p_t^*)$$

where $ppp = (p_t - p_t^* - s_t)$. If the PPP restriction (1,-1,-1) had been acceptable in all cointegrating relations in the I(2) model, the VAR model analysis of this transformation would have been empirically equivalent and no long-run information would have been lost by the transformation.[30] However, based on the results reported in the previous section, the joint restrictions were not data consistent in I(2) data set and some information is lost by using transformed vector. Thus, we have to admit that the real exchange rate transformation is a little problematic when implied restrictions are not statistically acceptable. Another solution would have been to remove I(2) by differencing. This would have resulted a greater information loss and would be inconsistent with economic theory. Since the likelihood function of the real exchange rate transformed model is now changed as compared to the analysis in a previous section, we have recalculated the misspecification tests. The stochastic specification regarding residual correlation, heteroscedastisity and normality indicates that the model can be considered as a satisfactory description of the data generating process.

Under the assumption that the model is a well defined statistical process, we next determine the number of cointegration vectors. The status of deterministic terms like constant or trend has been determined testing the joint hypothesis of both the rank order and the deterministic components as suggested in Johansen (1992). Trends in the cointegration relations are supported by this test procedure.[31] The hypothesis of the exclusion of trend in a cointegration space is also rejected $\chi^2(5) = 14.48$ with the p-value 0.01.

Although test statistics accept the modelling of the vector where a linear trend is included in the cointegration space, this choice is partly problematic in this data vector. The linear trend should clearly be present in the I(2) model since prices have got a linear trend in them, but not necessarily

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[30] See the discussion in Juselius and Toro (1999) and Juselius and MacDonald (2000).

in this real exchange rate transformed I(1) model. If the trend is excluded in the cointegration space, however, the largest estimated eigenvalue of the companion matrix is outside the unit circle.[32] Under the assumption of the cointegrated VAR model, the eigenvalues should be inside the unit circle or equal to unity. Eigenvalues outside the unit circle correspond to explosive processes and the model we have chosen is not an adequate description of such data. Thus, we have chosen the model with the linear trend in cointegration space although admitting this choice is partly problematic.

It is a common practice to use the trace test to determine the number of cointegration vectors. However, the trace test has a low power against near cointegration alternatives especially in a small sample. Beside this, it is not straightforward to use the tabulated critical values for the trace test since the distribution of these are simulated under the assumption no weak exogeneity.[33] Because we have found evidence on this phenomenon no conclusion will be solely based on trace test result but, in addition, the inference of numbers of cointegration vectors will be defined detecting the roots of characteristic polynomial. This test procedures supports the choice of r =2 which is in line with the pervious findings in I(2) model. The signs of I(2) components have now disappeared, i.e. the roots of the characteristic polynomial are now consistent with I(1) model.[34]

To investigate the time series properties of the individual variables and their status in the system, three different tests are reported in Table 4.5. The test of stationarity indicates that none of the variables can be considered stationary. Based on the test for long-run exclusion all variables are found to be significant for the long-run structure.[35] The test of long run weak exogeneity investigates the absence of long run levels feedback. This test shows that the real exchange rate and German productivity can both be considered weakly exogenous for β. The test of both being jointly weakly exogenous is accepted based on $\chi^2(4) = 0,84$. The data strongly supports the hypothesis that the real exchange rate variable is not an adjusting variable in the cointegration space. This finding is not conditional on the number of cointegration vectors. Thus, the test of weak exogeneity provides evidence which is not in favor of our theory for determination of the real exchange rate.

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[31]We have to consider these test results with great caution. There is a small sample and dummy problem in the asymptotic tables used in this test procedure.

[32] This finding is not sensitive for lag structure. Similarly no simple dummy structure changes the explosive property of the model if constant is included in the cointegration vector.

[33] See Jore *et al.* (1993).

[34] Five largest roots of the companion matrix are (0,95;0,95;0,88;0,47;0,47) in unrestricted model and (1,00;1,00;1,00;0,63;0,47) in restricted (r=2) model.

[35] Although not reported here the test for long-run exclusion accepts the linear trend in the cointegration space.

Table 4.5. Properties of system variables.

	Chi(2)	ppp	dp	dp*	Pro	Pro*
Stationarity	5,99	36,83	22,20	45,66	23,13	30,86
Exclusion	9,49	30,66	25,11	23,83	35,97	35,83
Weak exogeneity	5,99	0,46	25,16	49,29	0,47	14,93

Although the test of weak exogeneity suggests the Harrod-Balassa-Samuelson theory is not an adequate theory to explain the trending behavior of real exchange rate, we have tested whether there is a long-run equilibrium among the variables determined by this theory. We cannot find any evidence for the Harrod-Balassa-Samuelson effect using this data vector. Thus we can conclude that the real exchange rate is not an adjusting variable and there is no long-run relationship between productivity variables and the real exchange rate, i.e. the productivity differential does not seem to be able to explain the nonstationary behavior of the real exchange rate.

5. STOCHASTIC FORMULATION OF THE REAL EXCHANGE MODEL

There has recently been an increased interest in the small sample properties of cointegration tests.[36] Since the null hypothesis of a unit root is not necessarily reasonable from an economic point of view, the low power is a serious problem. Economic theory suggests often *a priori* hypothesis for the number of independent trends. This is a strong argument for building the choice of r on economic theory as well as on statistical information in the data. In this section we will give a stochastic formulation of the real exchange rate model.

We will now analyze the econometric consequences of theoretical real exchange rate models. PPP and other relations will be given a stochastic formulation based on the decomposition of the data into once or twice cumulated shocks and a stationary component. In order to illustrate the main target of this section we first examine the Harrod-Balassa-Samuelson real exchange rate model in a world with no market rigidities, no trade barriers, no restrictions on capital movements, no transportation costs and fully integrated goods markets. Based on the results presented in the previous section we assume that prices are I(2) in both countries. This is again only the specific assumption for this time period not a general assumption on the time series properties of price indices. The data generating process (DGP) could then be represented as:

$$
\begin{bmatrix} p_t \\ p_t^* \\ s_t \\ \Pr o_t \\ \Pr o_t^* \\ \Delta p_t \\ \Delta p_t^* \end{bmatrix} = \begin{bmatrix} 1 \\ 1 \\ 0 \\ 0 \\ 0 \\ 0 \\ 0 \end{bmatrix} \begin{bmatrix} \Sigma\Sigma \mu_{1i} \end{bmatrix} + \begin{bmatrix} 1 & 1 \\ 1 & 1 \\ 0 & 0 \\ 0 & 1 \\ 0 & 1 \\ 1 & 1 \\ 1 & 1 \end{bmatrix} \begin{bmatrix} \Sigma \mu_{1i} \\ \Sigma \mu_{2i} \end{bmatrix} + X_0
$$

where p_t is home country price index, p_t^* foreign country price index, s_t is nominal exchange rate, $\Pr o_t$ home country productivity, $\Pr o_t^*$ foreign country productivity. $\Sigma\Sigma$ indicates twice cumulated shock, Σ once cumulated shock and X_0 includes a stationary component and possibly also a deterministic trend. Under the assumption of flexible prices and strong market integration between Germany (home country) and the United States (foreign country) we would except one common nominal price trend, μ_{1i}, in the data. The other shock of our model, $\Sigma\mu_{2i}$, is defined as a real shock.

The price differential between economies is a stationary relation due to the strong market integration. Thus, there is a short run price homogeneity between Δp_t and Δp_t^* which makes p_t and p_t^* CI(2,0) with cointegration vector (1,1). The real exchange rate is a stationary relation between the price differential and the stationary nominal exchange rate. It is also plausible to assume in this complete market world that there are rapid technology transfers between countries, i.e. there are identical technologies in both countries. This creates a stationarity relation between productivity variables. Thus, the number of cointegrated vectors is three (one extra stationary vector created by the stationary nominal exchange rate variable). There is also one I(2) trend and one I(1) trend.

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[36] See, for example, Johansen (1998).

5.1. DECOMPOSITION OF THE PRICES

There is no empirical support in our data for a stationary relation between prices implying a lack of complete market integration between two goods markets. This is in line with the discussion in Goldberg and Knetter (1997). According to Goldberg and Knetter (1997) there appears to be price discrimination between European and American markets, dictated by distinct competitive conditions in those markets. In fact, our finding is $p_t - p_t^* \sim I(2)$ and there seems to be no short run price homogeneity from which follows $\Delta p_t - \Delta p_t^* \sim I(1)$. Because even the inflation differential seems to be a non-stationary variable there might be a structural change in inflation process especially in the US as seen in the Figure 8 (Appendix 2). Figure 5.1 demonstrates the nonstationarity of the inflation rate spread. [37]
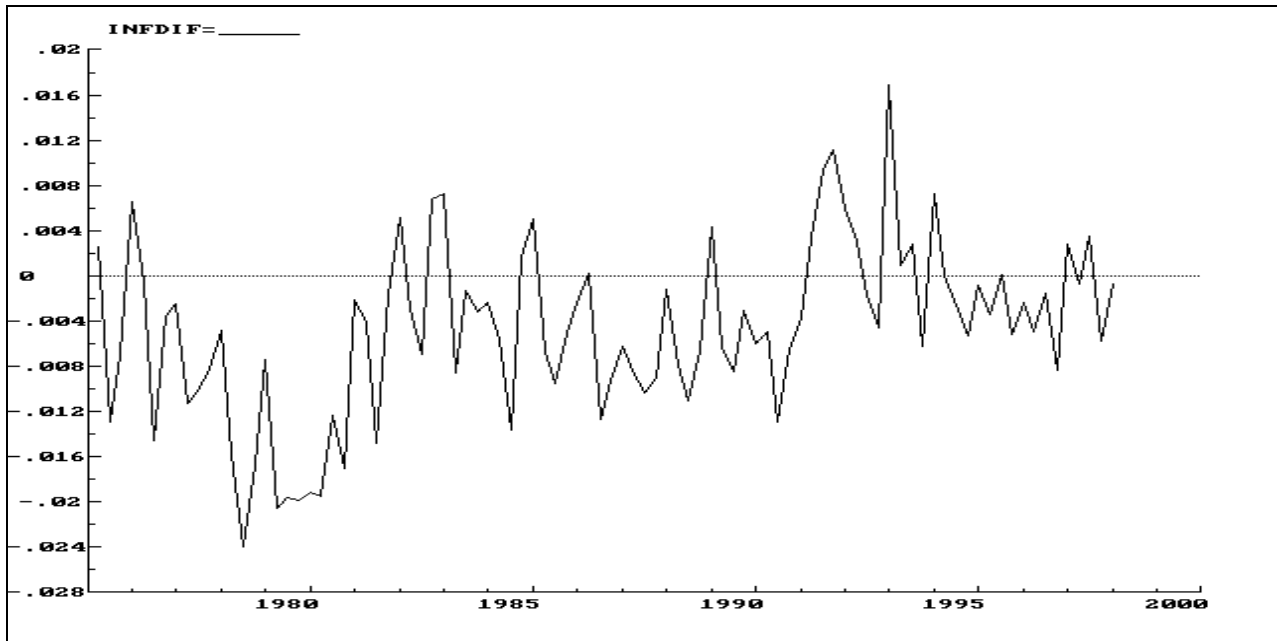

Figure 5.1. Inflation spread.

Relaxing the assumptions of perfect market integration and flexible prices, thereby allowing for different trends in nominal prices would decrease the amount of cointegrated vectors between two countries. The trend components of two price indices can be written as:

$$p_t = a_{11} \sum\sum \mu_{1i} + a_{12} \sum\sum \mu_{2i} + b_{11} \sum \mu_{1i} + b_{12} \sum \mu_{2i} + X_0$$

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[37] See also the results presented in Table 6.3.

$$p_t^* = a_{21} \sum \sum \mu_{1i} + a_{22} \sum \sum \mu_{1i} + b_{21} \sum \mu_{1i} + b_{22} \sum \mu_{2i} + X_0 \ ,$$

where $a_{ii} \neq 0$ (i = 1,2) indicating that prices are I(2). Based on the results in I(2) analysis, prices seem to include two I(2) trends. Following the above statements we define them as nominal and real shocks.[38] In our data set coefficients (a and b) are not necessary zero or one as in an illustrative example. We also assume, $a_{11} + a_{12} \neq a_{21} + a_{22}$, which would be consistent with the different stochastic price trends, excluding the possibility of a stable cointegration relation between two price indices.

5.2 DECOMPOSITION OF THE NOMINAL EXCHANGE RATE

In order to discuss the real exchange rate we also must define the statistical properties of nominal exchange rate. The results of I(2) analysis show that the nominal exchange rate might be affected by one I(2) trend but only very weakly. However, based on strong theoretical arguments, the nominal exchange rate and price differential should, *a priori,* share a common trend. This is also supported by the finding that the price differential and the nominal exchange rate are cointegrated (2,1). Thus, we conclude that there is a long-run relation between these variables but the statistically necessary I(2) trend in the nominal exchange rate, which is very difficult to identify empirically, partly reflects a weak and unstable relation between the nominal exchange rate and price differentials.

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[38] Note that there are two time periods during which the relation between two price indices is especially weak. The U.S. inflation decreased more than the German inflation during the first half of eighties and German reunification increased the German inflation substantially in the beginning of eighties.
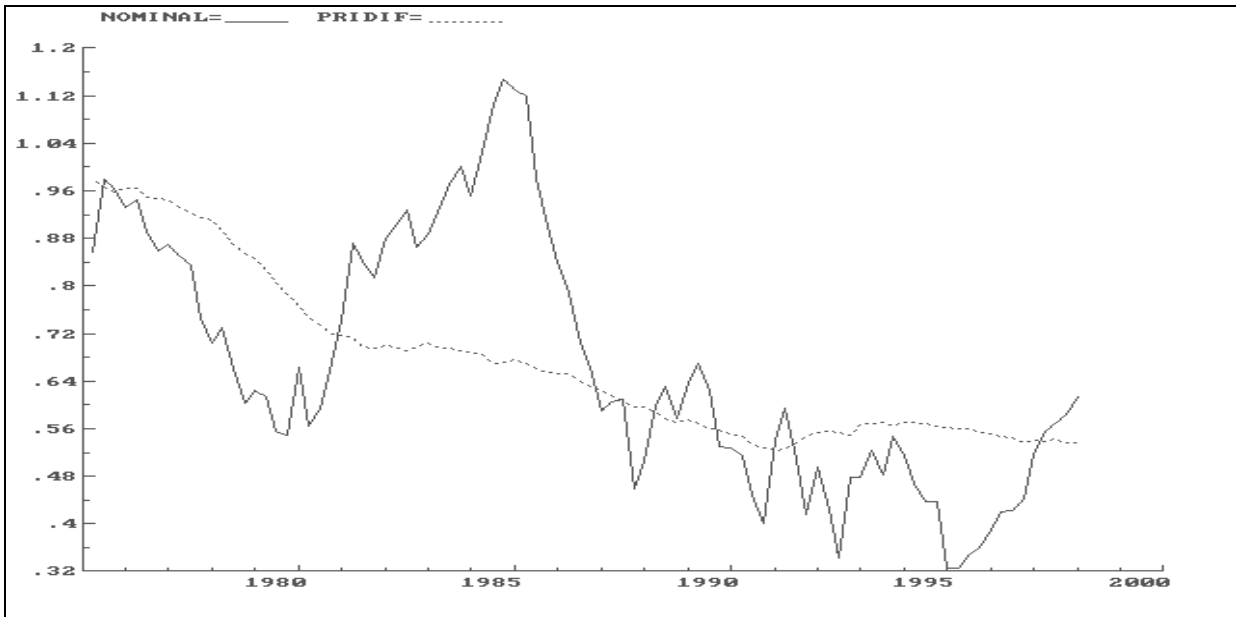
Figure 5.2. Nominal exchange rate and price differential.

The large deviations from the long-run price trend of nominal exchange rate in the chosen data are shown in Figure 5.2 above. This together with the econometric evidence of the I(1) analysis suggests that the nominal exchange rate has also been strongly affected by an additional stochastic I(1) trend, $\sum \mu_{3i}$, which does not seem to affect the price differential.

Juselius and MacDonald (2000) identified this extra stochastic trend as a "reserve currency" trend. The role of the U.S dollar as a reserve currency is likely to have resulted in permanent shocks to the nominal exchange rate that are unrelated to pure price shocks. However, we have modeled the third stochastic shock explicitly by using a real oil price shock as a proxy for permanent shock in the U.S dollar. Following the results presented in the next section higher oil price leads to the appreciation of the U.S. dollar in the long run. Thus, in order to explain a nonstationary real exchange rate in our data set, we should think of the oil price shocks as given rise to the dollar appreciation relative to the German mark which, in turn, creates a persistent price differential between the two countries.

The positive relationship between the U.S. dollar and oil price is partly problematic because, being a major importer of crude oil, higher oil price worsen the U.S. terms of trade.[39] This should depreciate the U.S. dollar not appreciate it relative to the German mark because oil price changes on

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[39] Crude oil is a complement to home goods in a medium run. In the long run Germany has decreased its dependence on oil more rapidly than the United States. See Backus and Crucini (1998).

the United States terms of trade affect more negatively than German terms of trade.[40] Interestingly, Amano and van Norden (1995), present evidence of a similar effect for Canada, where higher oil price leads to the weaker Canadian dollar relative to the U.S. dollar despite the fact that Canada is a substantial exporter of oil and the U.S. is a net importer of crude oil.

However, this finding is not entirely counterintuitive. First, the save haven effect might appreciate the U.S. dollar, because of increased uncertainty in a world economy created by the oil price shock.[41] Secondly, according to Krugman (1983), the effect of the price of oil depends on whether the burden to a country to country's balance of payments created by higher oil imports is greater or less than the improvement due to OPEC investments and imports. Thus, it is important to consider the effect of oil price shocks on exchanges rates in a multi country framework.

In a three-country world (Germany, the United States and OPEC) higher oil price will transfer wealth from the oil importers (Germany and the United States) to the oil exporters. While American and German current accounts are thus worsened, however, there is an improvement in capital accounts as OPEC invests its trade surplus in dollars and marks. Whether the net effect is favorable or unfavorable for the dollar depends on whether OPEC investment in dollars is more or less than America's share of the industrial world's current account deficit.

If only the capital account is considered, the implicit assumption is that the OPEC spending lags behind income. Thus, over time the balance of payments effect of higher oil prices depend upon its preferences for goods, i.e. trade flows are more important in the long-run. According to McGuirk (1983) the net trade effect is also positive. The net improvement in the U.S. balance of trade would then also require a real appreciation of the U.S. dollar.

Following the discussion on temporary appreciation in the section 2, the uncertainty of the permanence of the nominal exchange rate changes will widen the range within which the price differentials can fluctuate. If the uncertainty as to the permanence of the shock causing nominal exchange rate changes is high, the risk adjusted profit of arbitrage of goods is not necessarily high enough for arbitrageurs to engage in, although the price differential would be substantial. If the oil

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[40] See the discussion in Backus and Crucini (1998).
[41] In fact, this explanation is very similar with the "reserve currency" explanation discussed in Juselius and MacDonald (2000).

price is the source of large and temporary changes in the nominal exchange, there is an incomplete price arbitrage of tradable goods after the nominal exchange rate shock.

Following the above discussion the nominal exchange rate is now defined as:

$$s_t = a_{31} \sum \sum \mu_{1i} + b_{31} \sum \mu_{1i} + b_{32} \sum \mu_{2i} + b_{33} \sum \mu_{3i} + X_0$$

Again, if $a_{31} \neq 0$, the nominal exchange rate is I(2). There is weak evidence only for one I(2) trend in the decomposition of the nominal exchange rate presented in section 4. This is not, however, the basic reason for long-run deviations from price differential trend, because the nominal exchange rate and the price differential are cointegrated CI(2,1). In order to explain these large and persistent deviations from PPP, there should exist also the third once cumulated shock in the nominal exchange rate decomposition as discussed above.

5.3. DECOMPOSTION OF THE FULL MODEL

The above discussion is summarized in the following matrix:

$$
\begin{bmatrix}
p_t \\
p_t^* \\
s_t \\
\Pr o_t \\
\Pr o_t^* \\
Oilp_t \\
\Delta p_t \\
\Delta p_t^* \\
\Delta s_t
\end{bmatrix}
=
\begin{bmatrix}
a_{11} & a_{12} \\
a_{21} & a_{22} \\
a_{31} & 0 \\
0 & 0 \\
0 & 0 \\
0 & 0 \\
0 & 0 \\
0 & 0 \\
0 & 0
\end{bmatrix}
\begin{bmatrix}
\sum \sum \mu_{1i} \\
\sum \sum \mu_{2i}
\end{bmatrix}
+
\begin{bmatrix}
b_{11} & b_{12} & 0 \\
b_{21} & b_{22} & 0 \\
b_{31} & b_{32} & b_{33} \\
0 & b_{42} & 0 \\
0 & b_{52} & 0 \\
0 & 0 & b_{63} \\
a_{11} & a_{12} & 0 \\
a_{21} & a_{22} & 0 \\
a_{31} & 0 & 0
\end{bmatrix}
\begin{bmatrix}
\sum \mu_{1i} \\
\sum \mu_{2i} \\
\sum \mu_{3i}
\end{bmatrix}
+ X_0
$$

The cointegration properties of the data can be discussed using the above matrix. If $a_{12} = a_{22}$ and $a_{11} - a_{21} = a_{31}$ then $p_t - p_t^* - s_t = (b_{11} - b_{21} - b_{31})\mu_{1i} + (b_{12} - b_{22} - b_{32})\mu_{2i} + X_0$ is at most I(1). In our data it is also at least I(1) because there is the third I(1) shock, $\Sigma \mu_{3i}$, which does not directly

affect price differential, indicating that the real exchange rate is a nonstationary variable. Thus, it is not possible to find a stationary relationship between the classical HBS variables without modeling the third shock since oil price affects through the nominal exchange rate. Because we model the third shock explicitly it will increase the number of variables by one, but it will also increase the number of cointegrated variables at least by one. Thus, the number of cointegration vectors in the full model should be at least three. (r = 2 in a previous analysis).

6. THE REAL EXCHANGE RATE MODEL WITH OIL PRICE

In this section we will discuss estimation results based on the data vector including real exchange rate, two productivity variables, price differentials and a linear trend as in a previous estimation, but now also a real oil price variable. We will show the importance of real oil price for determination of the stationary real exchange rate. The dummy vector now includes also a special dummy designed for the Gulf War. The chosen VAR model needed the following three dummy variables:

$$D_t^{'} = D91, D91, D93,$$

where D91 and D93 are the same dummies as those discussed in section 4. D91 is a Gulf War dummy measuring a transitory shock defined as plus one in 1990:3, minus one 1991:1 and zero otherwise. Despite the inclusion of this dummy, the residual of oil price variable is not normally distributed and also the multivariate normality assumption is violated. This is not surprising since the real oil price variable was especially chosen to explain the variation in the real exchange rate but not vice versa, meaning that the selected variable set is probably not sufficient to account for the variation in the real oil price. The real oil price variable might be weakly exogenous for the long-run parameters of interest, which would make the deviation from normality less important (see test results in Table 6.2.). Residual normality is also mainly rejected due to excess kurtosis. Because cointegration results appear robust to excess kurtosis, we have ignored this normality problem. The results of the misspecification tests are reported in Table 6.1, where a significant test statistic is given in bold face.

Table 6.1. Misspecification tests and characteristic roots

Multivariate tests					
Residual autocorr.	LM(1)	CHISQ(25) =	44,95	p-val =	0,16
	LM(4)	CHISQ(25) =	40,18	p-val =	0,29
Normality	LM	CHISQ(10) =	27,98	p-val =	0,01

Univariate tests	dppp	ddp	ddp*	dPro	dPro*	dOilpr
ARCH(2)	0,60	0,20	5,78	0,19	2,03	0,34
Normality	3,34	4,09	0,69	1,39	2,75	**23,74**
Skewness	0,38	0,40	-0,05	-0,01	-0,32	-0,79
Ex. Kurtosis	0,54	0,70	0,10	-0,68	-0,39	3,35
R-squared	0,36	0,76	0,63	0,34	0,43	0,54

Six largest roots of the process						
Unrestricted	0,93	0,93	0,89	0,57	0,57	0,52
r = 4	1,00	1,00	0,88	0,88	0,51	0,51
r = 3	1,00	1,00	1,00	0,66	0,66	0,50

The inference on the number of cointegrated vectors was again made on the basis of the roots of characteristic polynomial. This supports the choice of r =3, which is also consistent with the theoretical discussion in a previous section. As a sensitivity check, the roots under the choice r = 4 are also reported in Table 6.1. As a check of the properties of the system variables tests for stationarity, long-run exclusion and weak exogeneity are reported in Table 6.2.

Table 6.2. Properties of the variables if oil price is included in the data vector.

	Chi(2)	ppp	dp	dp*	Pro	Pro*	Oilpr
Stationarity	9,49	20,48	19,85	17,86	26,71	26,81	19,83
Exclusion	7,81	31,52	22,56	44,75	21,73	27,48	10,68
Weak exogeneity	7,81	11,00	26,11	57,61	1,33	17,80	7,93

All variables were found to be significant for the long run structure and the weak exogeneity results show that German productivity and possibly oil price can be considered weakly exogenous for the long term parameter β.[42] However, the most interesting result is that the real exchange rate is not weakly exogenous if the real oil price variable is included in the data vector.

6.1. STRUCTURAL HYPOTHESES ABOUT THE COINTEGRATION SPACE

The purpose of this section is to demonstrate that the some of the economic hypotheses discussed in section 2 and section 6 can be given a precise statistical formulation, and therefore, can be validly tested. In the ideal case the specification of the matrices should define a identifying structure with all freely estimated coefficients statistically significant and economically interpretable. Hypothesis 1-6 are of the form $\beta = \{H_i\phi_1, \psi_1, \psi_2\}$i=1..6; that is, they test whether a single restricted relation is in the cointegration space, leaving the other two relations unrestricted.[43] The results are shown in Table 6.3.

Table 6.3. Tests on the cointegration vectors.

	ppp	dp	dp*	Pro	Pro*	Oilpr	Trend	Chi(v)	p-value
H 1	0	1	-1	0	0	0	0	10,89(4)	0,02
H 2	0	1	-0,7	0	0	0	0	8,66(3)	0,03
H 3	0	0	0	1	-1	0	0	16,95(4)	0,00
H 4	0	0	0	1	-0,84	0	0	12,95(3)	0,00
H 5	1	0	0	0	0	0	-0,003	18,75(3)	0,00
H 6	1	0	0	0	0	0,53	0	12,88(3)	0,00
H 7	1	0	0	6,67	-6,67	0	0	13,31(3)	0,00
H 8	1	0	0	12,73	-11,75	0	0	11,63(2)	0,00
H 9	1	0	0	0	0	1,27	0,01	3,71(2)	0,16
H 10	1	0	0	-17,90	17,90	2,04	0	7,96(2)	0,02
H 11	1	0	0	-9,27	11,24	2,26	0	0,57(1)	0,45
H 12	-0,03	0	1	0	0	-0,02	0	0,06(2)	0,97
H 13	0	1	-1	-0,078	0	0	-0,001	0,36(2)	0,83

Hypotheses H(1) and H(2) are related to the inflation differential. A stationary relationship between inflation rates is a borderline case, being rejected at 5% but not at 1 % significant level. Hypotheses H(3) and H(4) are related to the long-run relationship between productivity variables in Germany and the United States. We have not found a stationary relation between the productivity variables in our data, which includes 23 years. This result may be an indication of an underlying catching up process, whereby the German productivity level has been converging to the corresponding US one. Thus, it seems to be taking a relatively long time period before the assumption of similar technology is accepted even in the case of two industrialized countries.

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[42] Again, a linear trend in the cointegration space was supported by this test procedure.
[43] For the derivation of the test procedures, see Johansen and Juselius (1992).

Hypotheses H5 and H6 are related to the real exchange rate variable. It is a common habit to model possible long-run deviations from PPP as linear trend, i.e. if long-run deviations are due to the Harrod-Balassa-Samuelson hypothesis the linear trend represents the productivity differential. Based on the results concerning hypothesis five we have not found evidence for this relation. H6 tests whether there is a stationary relation between the real exchange rate and the real oil price. Amano and Van Norden (1999) found this relation using the US real effective real exchange rate but here we have not been able to find a stationary relation between the real exchange rate and the real oil price.

Hypotheses H7-H11 test the real exchange rate relations in a multivariable environment. These are all tests of the type, $\beta = \{H_i \phi_i, \psi_i\}$ i= 7…11, where ($H_i \phi_i$)' is given by the hypothetical vector related hypotheses 7-11 in Table 6.3. and $\psi = (\psi_{i1}, \psi_{i2})$ is a matrix of unrestricted coefficients. Hypotheses H7 and H8 clearly reject the standard Harrod-Balassa-Samuelson hypothesis. H10 introduces the real oil price in a multivariable environment. The most interesting result is the hypothesis 11. We are able to find a traditional Harrod-Balassa-Samuelson result if the oil price variable is included in the data vector between the German mark and the U.S. dollar with the test statistic $\chi^2(1) = 0,57$.[44] The result suggests that the oil price may have been an important source of persistent real exchange rate shocks. However, before drawing any conclusions we have to first recognize the full cointegration space and also make a stability analysis.

H12 and H13 identified the other two cointegration vectors. H12 defines U.S inflation using the real exchange rate and oil price. H13 is probably best understood as an inflation differential between Germany and the U.S.A. The linear trend might indicate the increased anti-inflation credibility of the FED which has decreased the inflation differential between the currencies.[45] Undoubtedly, the economic rationality of these two relations is far from perfect. Thus, these relations give us reason to suspect that other variables than those "real economy" variables included in the analysis probably play an important role, and only with a much larger system it is possible obtain a model that satisfies both statistical and economic interpretability.

Using the above results it is now time to move a structural formulation of the full cointegration space expressed as the following joint hypothesis , $\beta = \{H_1 \varphi_1, H_2 \varphi_2, H_3 \varphi_3\}$, where the design

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[44] Typically the oil price effect has been analyzed using multilateral exchange rates. See MacDonald (1997) or Amano and Van Norden (1998).

matrix H_i defines the assumed structural representation.[46] The joint hypothesis about the long-run structure was formulated by the following design matrices:

$$
H1 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 0 \end{bmatrix} \quad H2 = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & 0 \end{bmatrix} \quad H2 = \begin{bmatrix} 0 & 0 & 0 \\ 1 & 0 & 0 \\ -1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}
$$

These matrixes combine H11, H12 and H13 and are together safely accepted with a test statistic 2,39 asymptotically distributed as $\chi^2(5)$ and p-value of 0,79. The estimates of the unrestricted β_{ij} coefficients and their asymptotic standard errors are given in Table 6.4.

Table 6.4. Cointegration vectors.

		PPP	Dp	Dp*	Pro	Pro*	Oilpr	Trend
Vector 1	Coefficient	1	0	0	-10,14	11,84	2,14	0
	Stand. error				-1,46	1,42	0,24	
Vector 2	Coefficient	-0,027	0	1	0	0	-0,02	0
	Stand. error	0,002					0,002	
Vector 3	Coefficient	0	1	-1	-0,092	0	0	-0,001
	Stand. error				0,02			0,000

The results in the above table show that all freely estimated coefficients are significant. All three cointegration relations are presented in Appendix 2.

It may be also informative to examine the structure of the α coefficients. The short run adjustment in the real exchange rate equation takes place primarily in inflation variables, but to some extent also in the real oil price variable. The real oil price is a borderline finding with the t-value 2,15.[47] Although it is possible that OPEC considers the effects of oil price changes, the oil price variable should probably be a weakly exogenous variable in this cointegration space. The real exchange rate, inflation and the U.S productivity variable are significant in the U.S inflation equation. The short

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[45] Obviously, the linear trend is designed just for this time period.
[46] This test procedure is discussed in Johansen and Juselius (1994).
[47] This partly explains the borderline finding of real oil price weak exogeneity in Table 6.2.

run adjustment to the third cointegration vector, defined as inflation differential equation, is mainly due to the US and German inflation.

The stability of the results has been checked by a recursive stability test suggested Hansen and Johansen (1993).
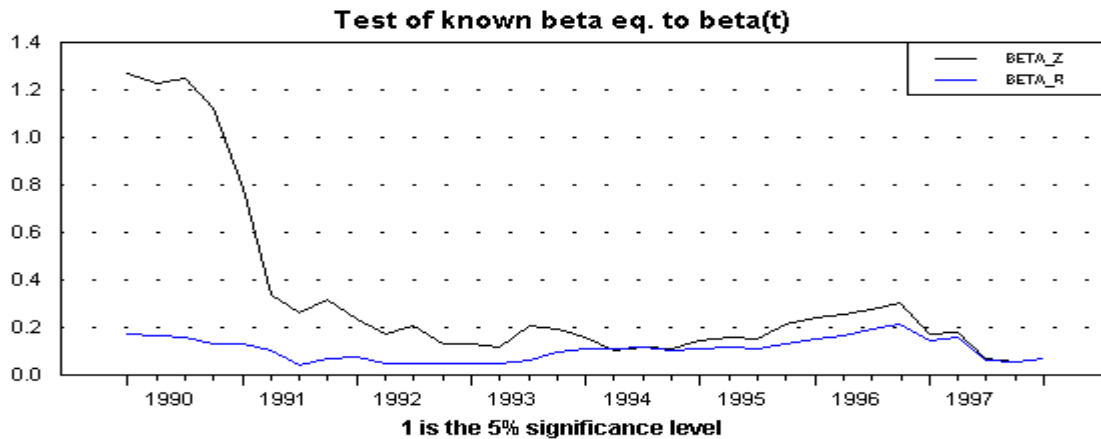
Figure 6.1.Recursive estimates (base year 1990).

Figure 6.1. shows recursively calculated test statistics for the test of a constant cointegration space. Beta Z (upper line), which shows actual deviations as a function of all the short run dynamics including seasonals and other dummy variables, displays stability of the cointegration space after an initial period of about one and half years. On the other hand, Beta R, which corrects for short run dynamics, clearly gives a stronger sense of stability of the cointegration space.[48]

Altogether, we find that the results are interpretable and provide insight into dynamics of the highly complex long-run adjustment process of the German mark-U.S. dollar real exchange rate. Generally results show that the real exchange rate and the goods market are related in such a way that the PPP hypothesis or the Harrod-Balassa-Samuelson effect alone cannot explain this complex structure.

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[48] We have tested the stability of the cointegration space using several base years. These findings also support the parameter constancy.

7. CONCLUSIONS

Recently, there have been arguments in the literature that the failure of PPP to hold even for traded goods may be largely an U.S. dollar phenomenon. Engel (1999) has shown that the movements in the relative price of traded goods are important for the U.S. real exchange rate. Canzoneri *et al.* (1999) use panel cointegration methods to examine the Harrod-Balassa-Samuelson effect. They argue that the problems with the Harrod-Balassa-Samuelson hypothesis lie in the failure of PPP to explain traded goods prices, especially for the U.S dollar.

We have discussed oil price as a possible source of the failure of PPP for traded goods, when the U.S. dollar real exchange rate is considered. As discussed in Froot and Klemperer (1989) exporters may respond differently to temporary and permanent changes in the exchange rate. When the dollar is temporary high, foreign firms will find investments in market share less attractive, and will prefer instead to let their current profits to increase. If oil price is the source of large and temporary changes in the nominal exchange, there is the incomplete price arbitrage of tradable goods after the nominal exchange rate shock.

The first model was the I(2) model. We have found no evidence for the stationary real exchange rate. In fact, the empirical evidence for the cointegration relation between price differential and the nominal exchange rate is ambiguous. I(2) property of the price differential may be due to structural changes especially in the U.S. prices. This was also confirmed by the findings of I(1) analysis, which showed the need for a linear trend in the cointegration vector understood as the inflation differential between Germany and the U.S.

The empirical analysis of the first I(1) data set was based on the real exchange rate transformed vector, where the real exchange rate is an I(1) variable. The joint restrictions were not data consistent in I(2) data set and some information is lost by moving to I(1). Thus, we have to admit that the transformation is problematic when implied restrictions are not statistically acceptable. Until we have full-fledged I(2) program there is not much we can do. However, I(2) analysis given in this study makes clear the importance of complete I(2) analysis in order to understand fully the relationship between the nominal exchange rate and prices.

A statistically acceptable I(1) model provides us the opportunity to examine the common stochastic trends between variables. Using the data vector with the real exchange rate variable we did not find

any evidence for the Harrod-Balassa-Samuelson hypothesis. Based on this result and the well known problems of the second component of Harrod-Balassa-Samuelson hypothesis (PPP for tradables), we extended our data vector with the oil price variable. Finally, we tested whether some of the basic economic hypotheses can be given a precise statistical formulation. Test results clearly show that we can accept the Harrod-Balassa-Samuelson hypothesis if the oil price variable is included in the data vector. This is consistent with the hypothesis of no-arbitrage condition if there is uncertainty as to the permanence of the shock causing relative tradable price changes. Similarly, we are not able find a stationary cointegration vector, if only oil price is included in the vector. Thus, we can conclude that both sources of non-stationarity have to be included in the estimation vector in order to find complete understanding of deviations from PPP.

Most of the time cointegration analysis rejects the hypothetical coefficients. This seems to imply that other variables than those included in the analysis might play a role and only with a much larger system it is possible to obtain a model that satisfies completely both statistical and economic interpretability. For example, as shown in Juselius and MacDonald (2000), interest rates are also very important for the real exchange rate determination. Thus, it would be interesting to extend our model with the asset market variables. This is especially important because the discussion of economic rationality of weak euro has focused mainly on interest rate differentials. The asset market approach together with the variables discussed in this study would give us more complete understanding of the equilibrium level to which the economic adjustment forces pull the exchange rate between the German mark and the U.S. dollar.

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APPENDIX 1

The data set used consists of quarterly time series observations from 1975:2to 1998:1 for Germany and the United States. The source of the data has mainly been the IMF International Financial Statistic (IFS). The productivity data is provided by the Bureau of Labor Statistics. Following variables are considered:

Nominal exchange rate is average-for-quarter observations defined as the German mark price of one US dollar.

Prices are consumer price indices for Germany and the U.S.

The average product of labor in manufacturing is used to measure **productivity**.[49] Unfortunately, the equivalent measure for the non-tradables sector is not available. Thus, the measure productivity may be appropriate only under the further assumption that trend movements in relative productivity in services are insignificant in the U.S.A. and Germany.

The choice of the average product of labor differs from most of the recent literature, which uses total factor productivity as a proxy for productivity.[50] This choice is not innocuous since labor shedding may introduce substantial differences between changes in average labor productivity and changes in total factor productivity. However, as Ganzoneri *et al.* (1999), have pointed out, the common habit of describing movements in total factor productivity using Solow residuals is problematic because interpreting movements in Solow residuals as exogenous supply shocks is far from perfect. Another important surplus of average labor productivity compared to total factor productivity is that it holds for a broader class of technologies than the Cobb–Douglas production function, which is used to compute Solow residuals.

To be precise, our measure of productivity is an index of labor productivity, constructed from real output per man hour in manufacturing in the U.S. and Germany. As discussed above, a drawback to this measure of productivity is that it does not control for differences in investment rates, but on the

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[49] Note that the construction of productivity series are not completely similar. Output series for Germany is based on value-added basis but the quarterly U.S. manufacturing output series is based on a sectoral output basis rather than a value-added basis. However, this should not prevent us to find empirically relevant relations between variables.
[50] See, for example, De Gregorio and Wolf (1994), De Gregorio *et al.* (1994), Strauss (1995) and Chinn and Johnston (1996).

other hand it avoids the many difficulties involved on measuring capital.[51] The comparisons of productivity data are limited to trend measures only; no reliable comparisons of levels of manufacturing productivity are not available.[52] Under the assumption that labor productivity in manufacturing reflects overall productivity in traded goods, we can assume that labor productivity estimates provide a less problematic measure of existing productivity differentials in tradables in the U.S. and Germany than total factor productivity estimates.

Oil price is the quarterly average of the spot price of oil in US dollars, deflated by the US CPI. Although we use the real oil price variable we cannot exclude the possibility that the construction of price indexes can affect on real exchange rate movements if oil price is weighted differently in the United States than in Germany. Suppose that

$$p_t^T = \theta oilp_t + (1-\theta)p_t^2 \text{ and}$$
$$p_t^{T*} = \pi oilp_t^* + (1-\pi)p_t^{2*}$$

where θ and π are weights in home and foreign country (foreign marked with *) price indicies, p_t^T is a price index of tradable goods, $oilp_t$ is an oil price index and p_t^2 non-oil price index. Lower case letters denote variables in logarithms. Even if the law of one price for each good held, if $\theta \neq \pi$, then $p_t^T - p_t^{T*}$ will change as $oilp_t$ moves relative to p_t^2. However, Engel (1999) recalculated several traded goods price indices (including Germany) using the U.S. weights and found only little real exchange rate effect from using different weights. Thus, oil price seems to affect real exchange rates by some other way than the different index weights alone.

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[52] Using average labor productivity we do not need data on sectoral capital stocks, which are likely to be quite unreliable.

[53] To compare manufacturing output across countries, a common unit of measurement would be needed, such as the U.S. dollar. Market exchange rates are not suitable as a basis for comparing output levels. What is needed is reliable PPP. Reasonably reliable PPP is available only for the total gross domestic product not for the manufacturing product.
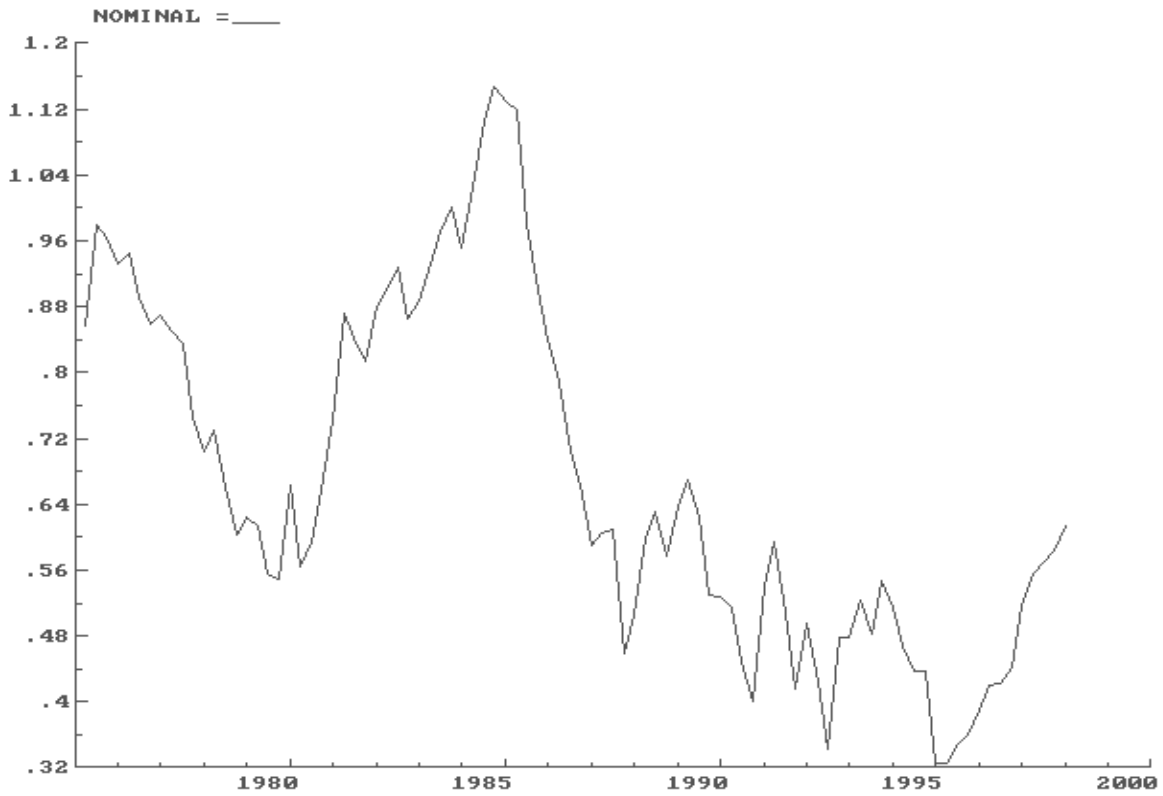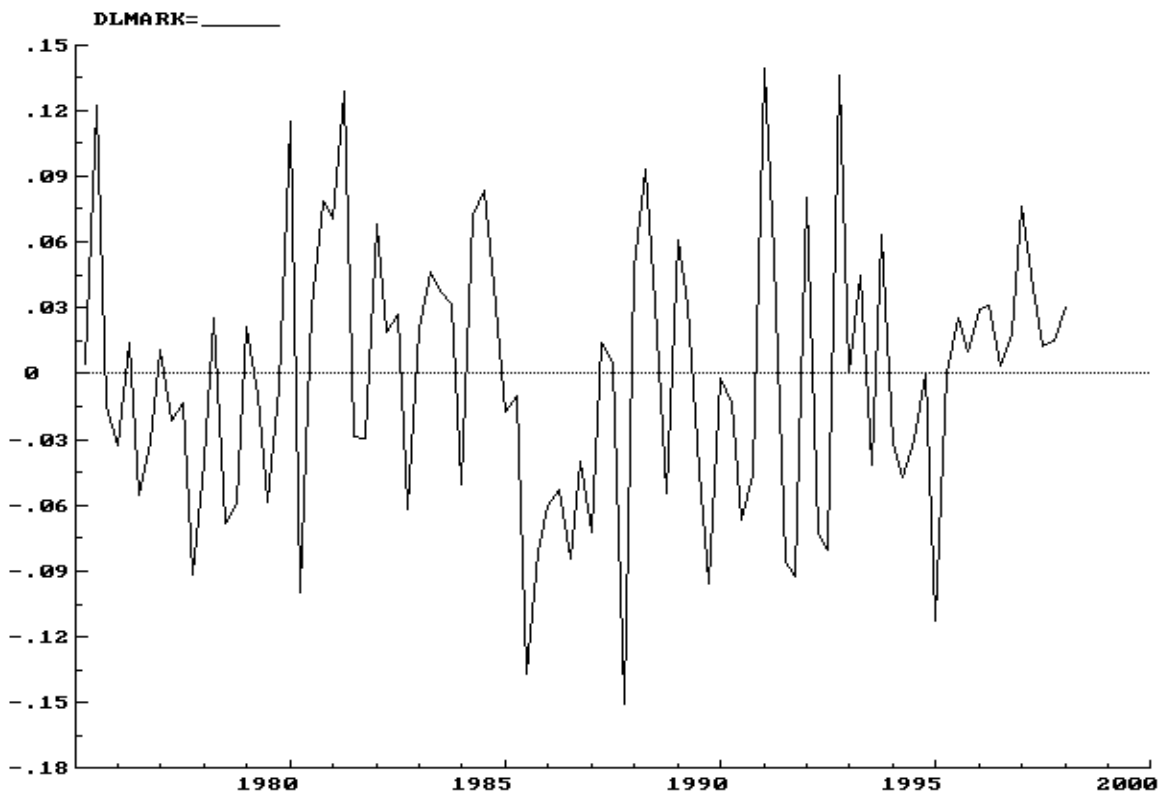
FIGURE 1. NOMINAL EXCHANGE RATE

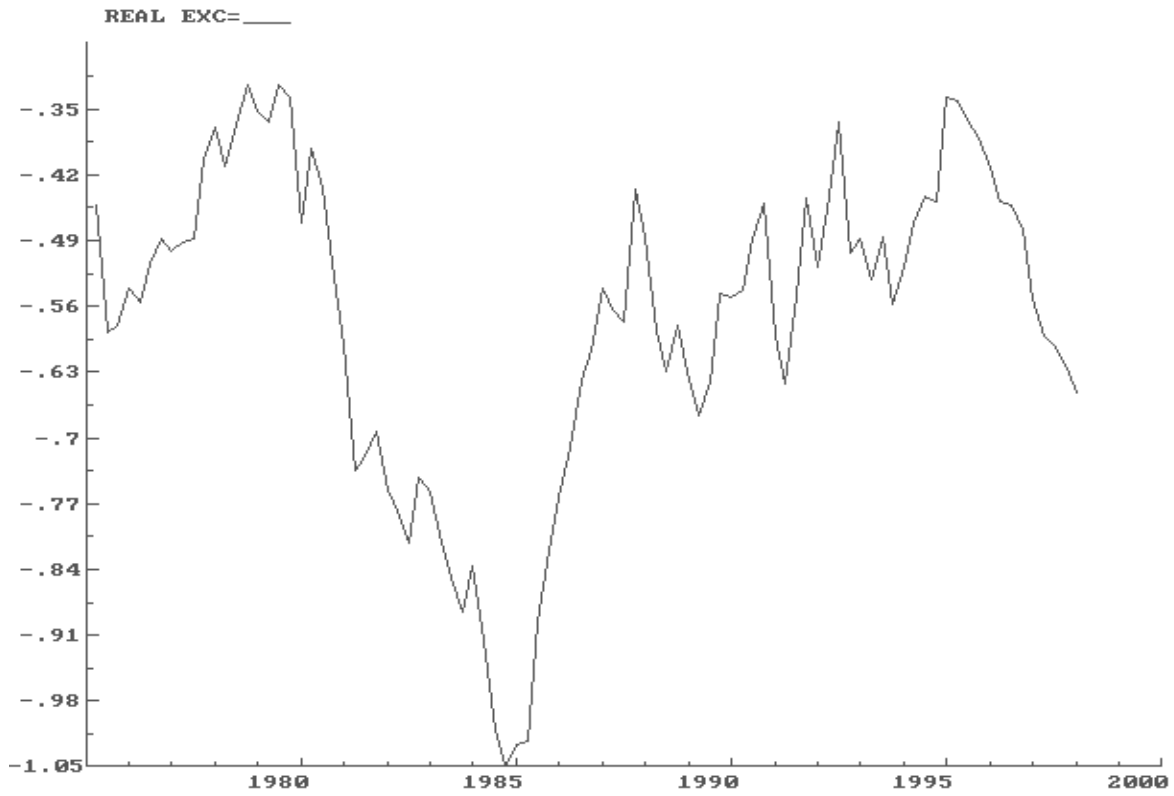FIGURE 2. DIFFERENCED NOMINAL EXCHANGE RATE
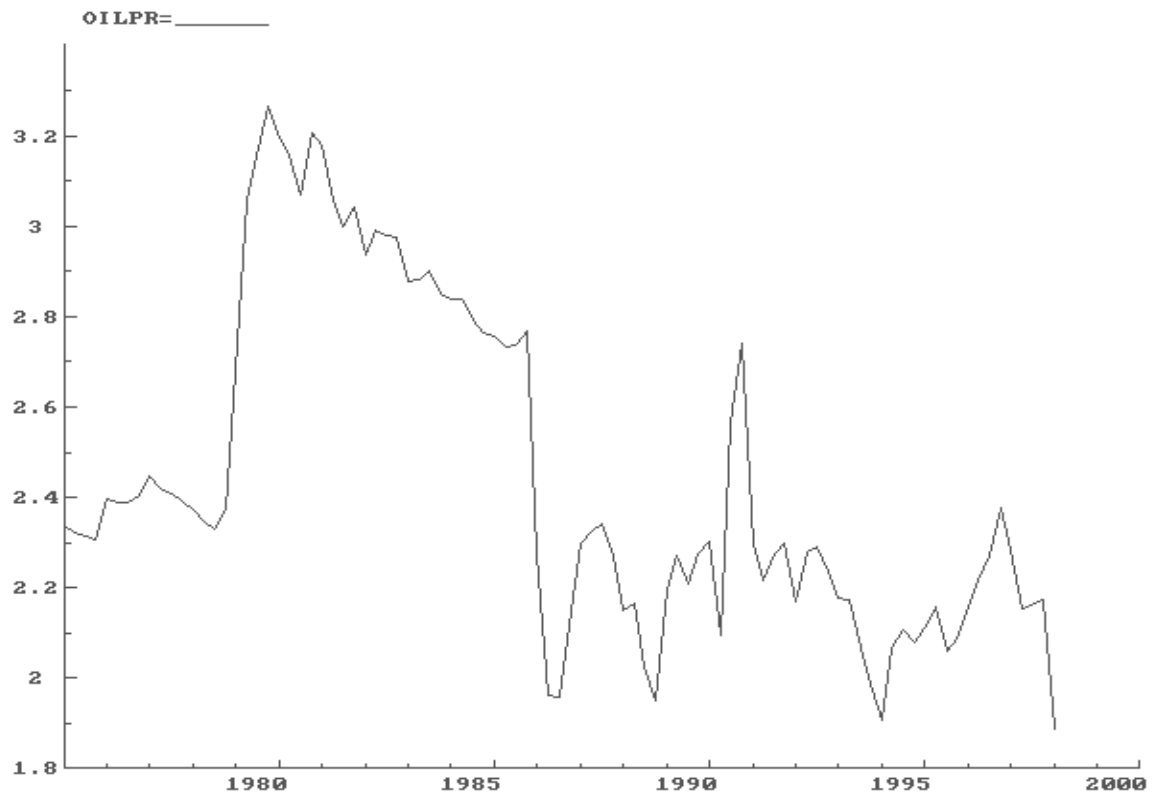
FIGURE 3. REAL EXCHANGE RATE
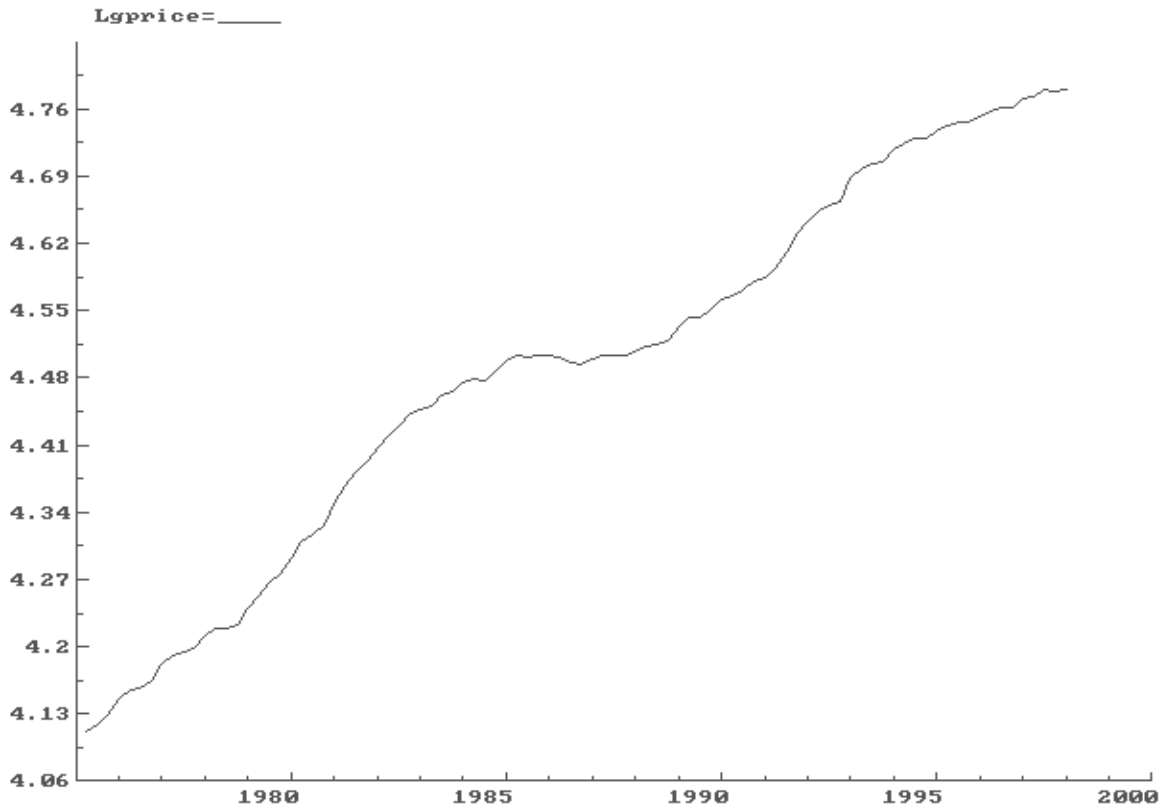

FIGURE 4. REAL OIL PRICE
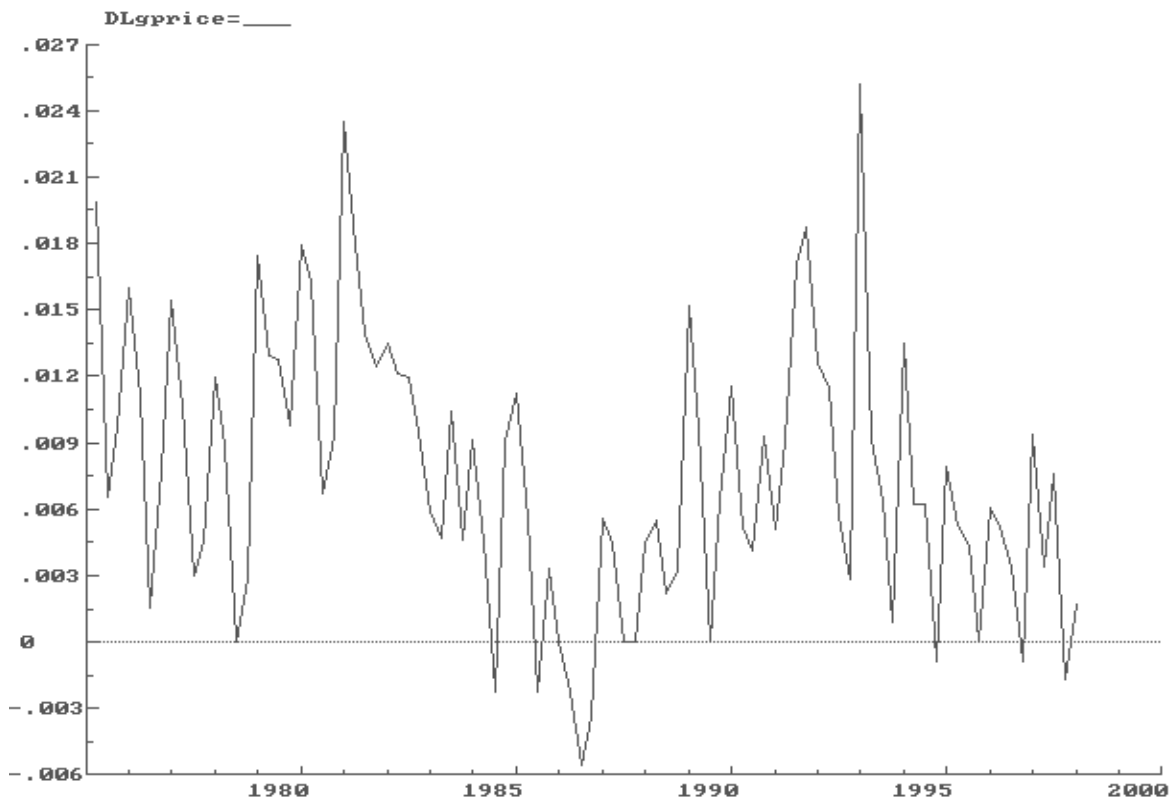
FIGURE 5. GERMAN PRICE INDEX
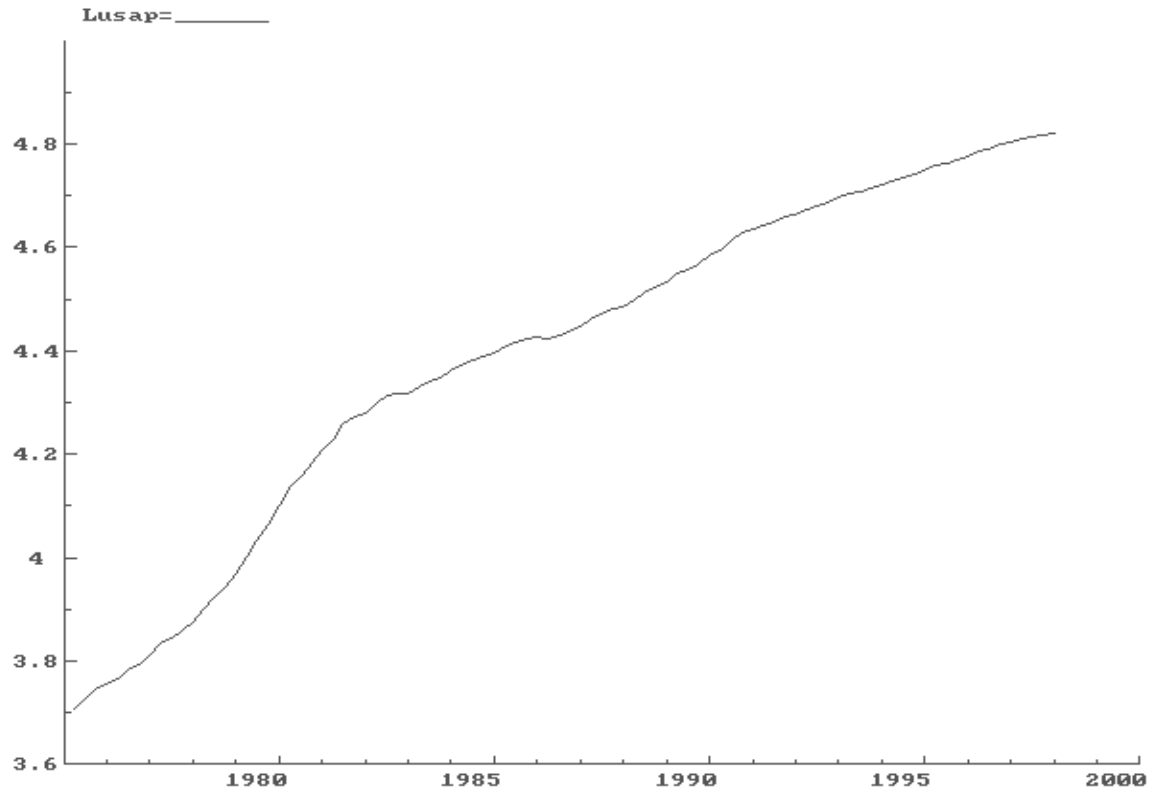

FIGURE 6. GERMAN INFLATION

FIGURE 7. U.S. PRICE INDEX
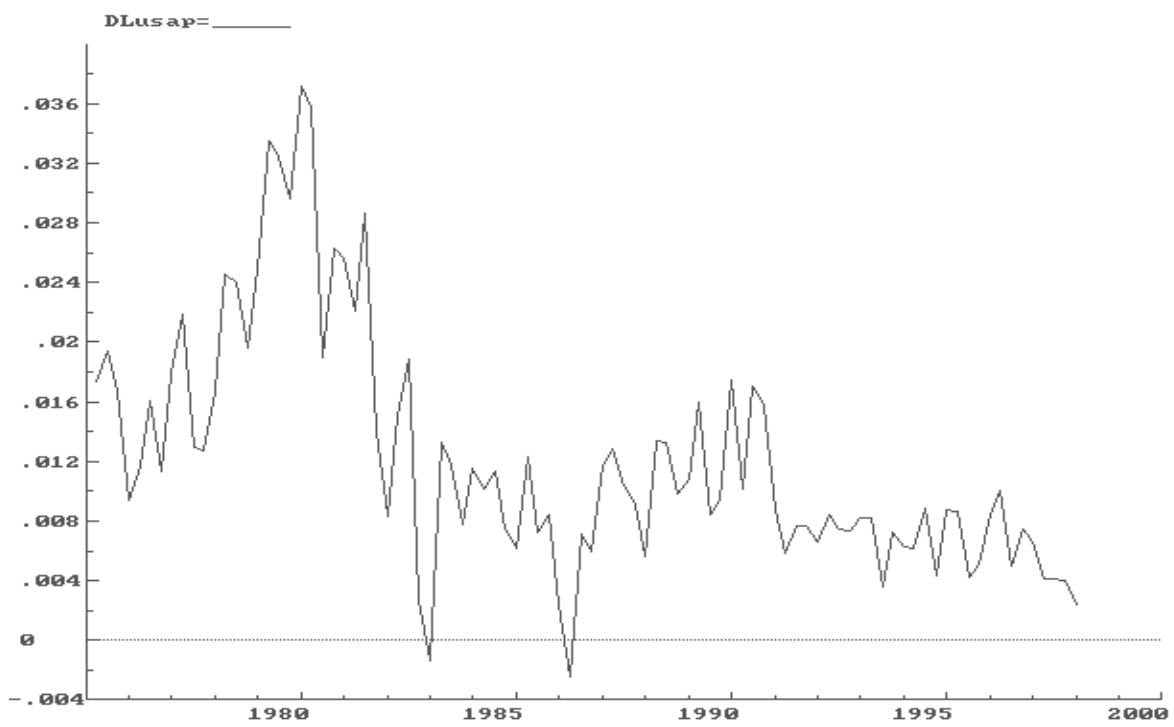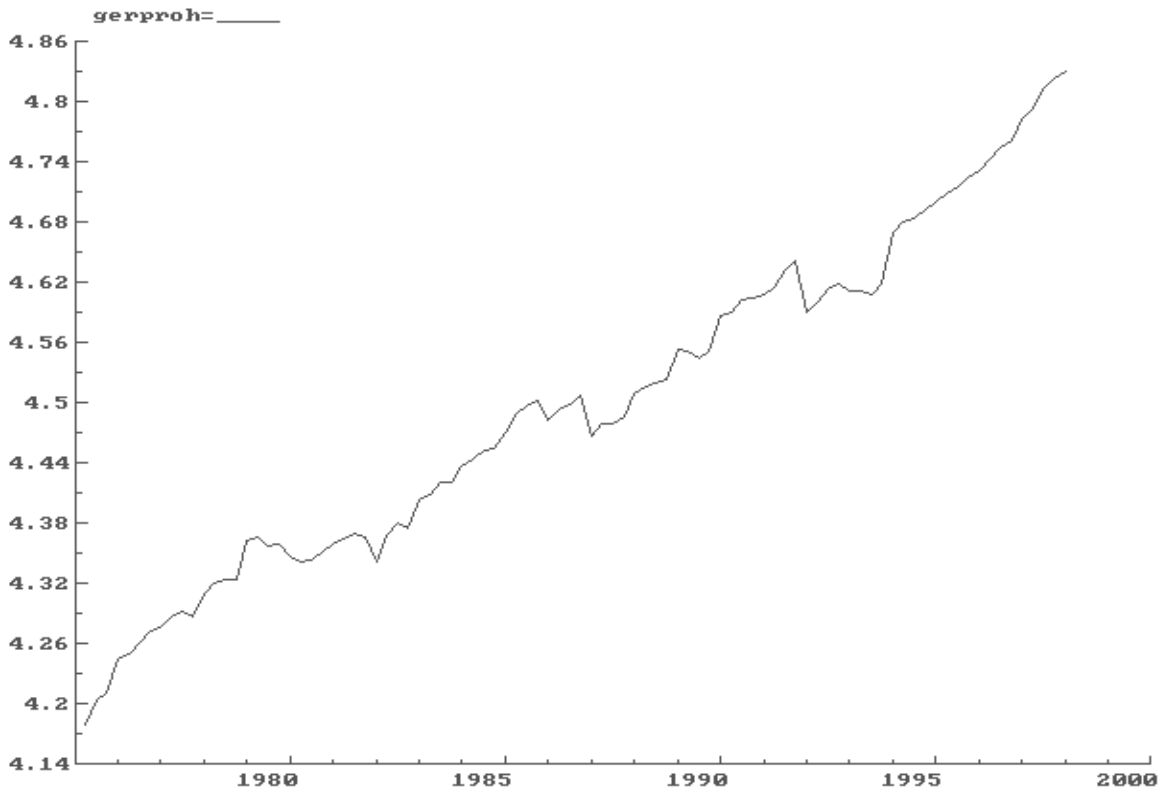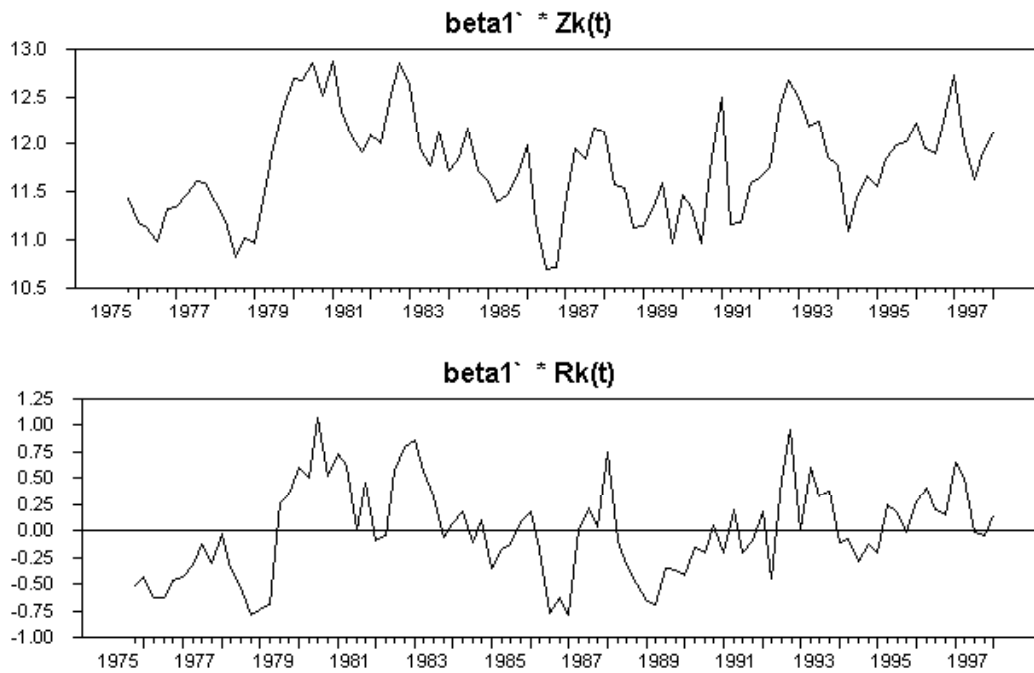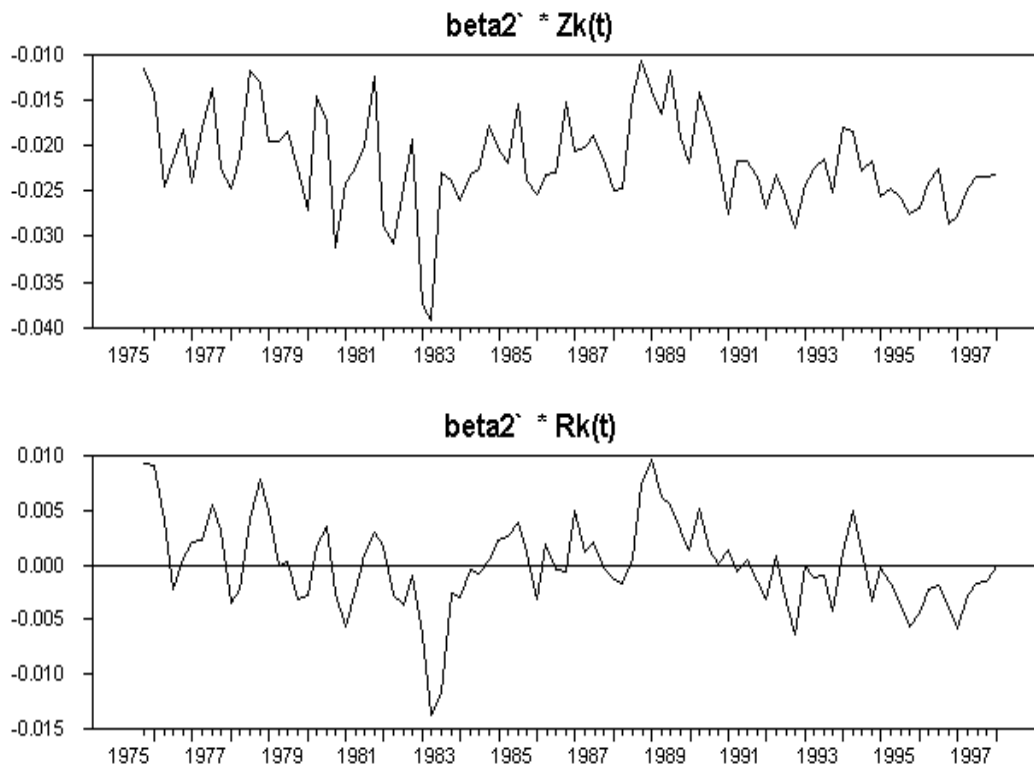

FIGURE 8. U.S. INFLATION
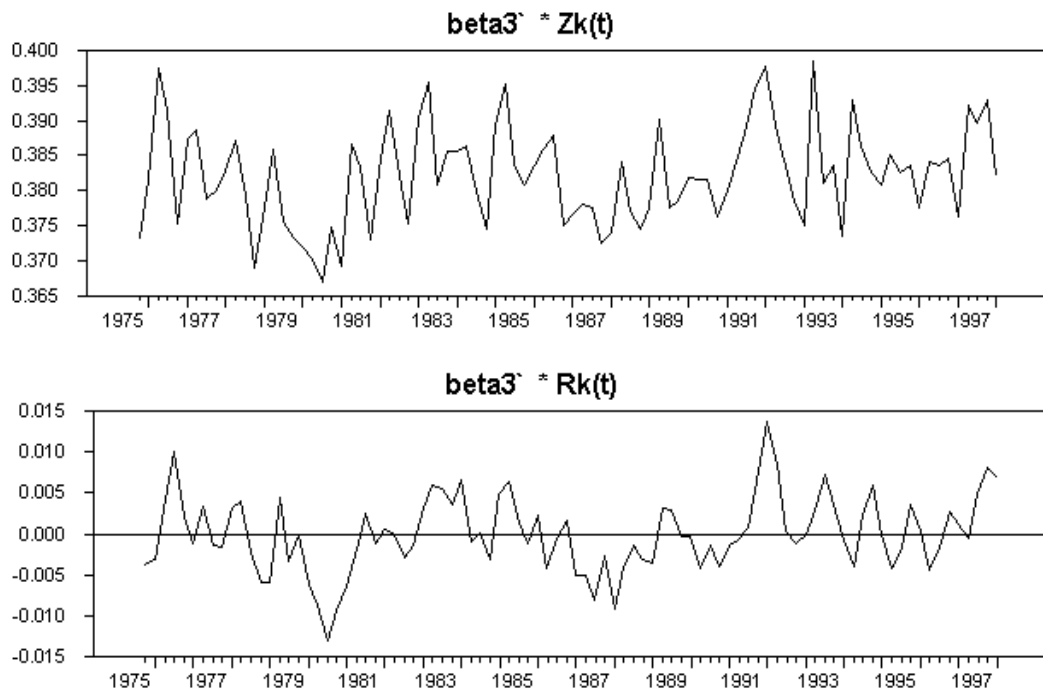
FIGURE 9. GERMAN PRODUCTIVITY
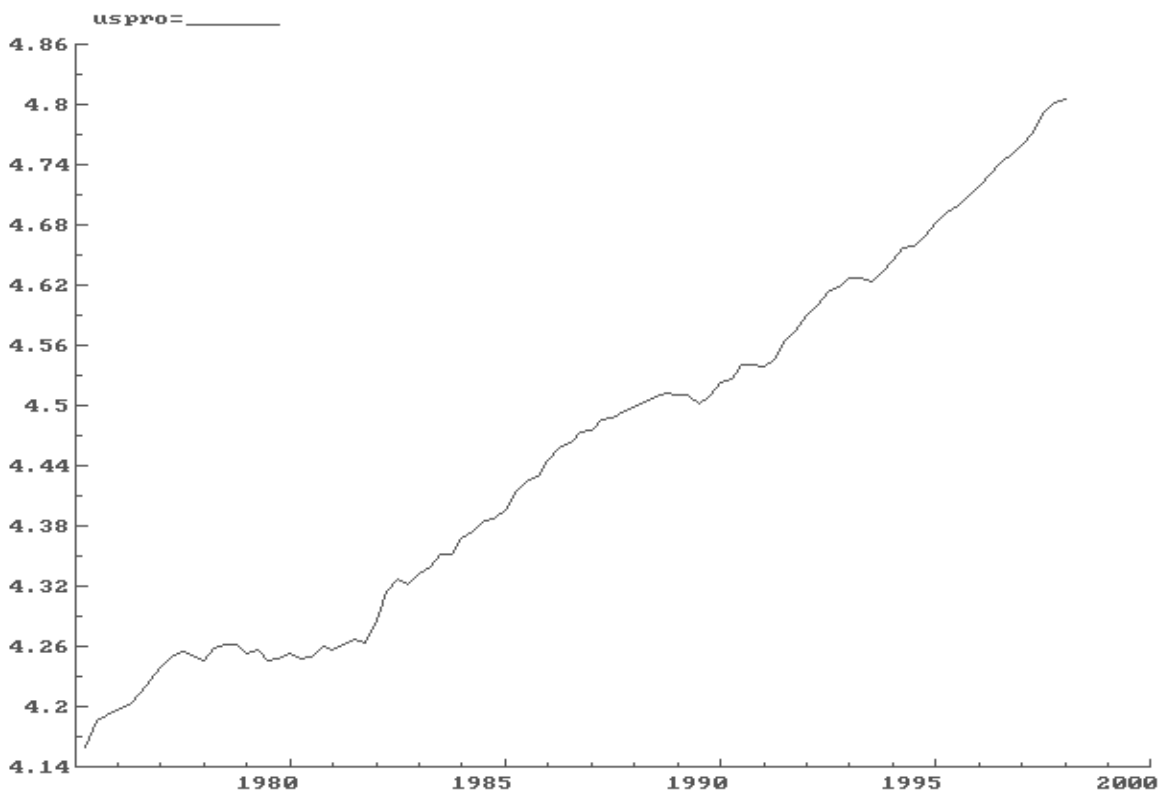
FIGURE 10. U.S. PRODUCTIVITY

APPENDIX 2

beta1` * Zk(t)

beta1` * Rk(t)

COINTEGRATION VECTOR 1

beta2` * Zk(t)

beta2` * Rk(t)

COINTEGRATION VECTOR 2

beta3` * Zk(t)

beta3` * Rk(t)

COINTEGRATION VECTOR 3

THE U.S. DOLLAR REAL EXCHANGE RATE

A REAL OPTIONS' APPROACH

Markus Lahtinen

Abstract

The aim of this paper is to discuss the determinants of the U.S. dollar real exchange rate fluctuation. We focus our analysis on the exchange rate effect on tradable prices. We explicitly consider the effects of profit maximizing foreign firms' entry decisions on the domestic tradable prices through the supply changes after a large appreciation. If firms face sunk entry costs when breaking into foreign markets, the extent of pass-through will depend on the expected changes of the exchange rate. Typically, exchange rate uncertainty is determined by the volatility of a continuous stochastic process. We extend the discussion also to consider possible jumps in the time path of the expected exchange rate. Finally, an interesting perspective is provided by a real option approach that emphasizes dynamic supply effects through sunk costs and uncertainty.

JEL classification: F21, F41

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1. INTRODUCTION

During some subperiods of floating rates traditional macroeconomic models, such as the monetary or portfolio model, explain monthly or quarterly exchange rate movements reasonably well, while during other subperiods their explanatory power completely disappears. If there were no relative structural shocks between two currency areas, a real exchange rate should be a stationary variable and it should follow the purchasing power parity hypothesis (hereafter PPP). However, the empirical evidence of large and persistent deviations from PPP is overwhelming. Strong evidence also suggests that the failures of the law of one price are not only significant, but that they also play a dominant role in the behavior of real exchange rate.[1]

A prominent question in this research is why the US real exchange rate exhibits large and persistent deviations from PPP. We discuss the determinants of the US dollar real exchange rate fluctuation since the collapse of the Bretton Woods system of fixed exchange rates in the early 1970's. The real exchange rate is measured using CPI deflators.

A number of studies have provided fairly convincing evidence that deviations from PPP derive largely from differences in relative traded goods prices across countries.[2] Recent studies have increased interest in the effect of fluctuations of the US dollar on the US tradables prices. Engel's (1999) results concerning the US dollar suggest that consumer prices for tradables goods behave in very much the same way as non-tradables consumer prices. These results are quite puzzling in view of the Harrod-Balassa-Samuelson hypothesis which holds that the relative tradables prices show little long-term variation across countries compared with the variation in relative non-tradables prices. Canzoneri *et al.* (1999) argue that the problems with the Harrod-Balassa-Samuelson hypothesis lie in the failure of PPP to explain the US dollar traded goods prices. The evidence above would suggest that the dominant component of a real exchange rate behavior is a nominal exchange rate even in the long run through the incomplete pass-through. Thus, we focus our analysis on a nominal exchange rate effect on tradable prices.

The incomplete pass-through of international price setting has been addressed in various ways in the literature. The most common analytical tool to examine incomplete pass-through has probably been the pricing-to-market approach, which presupposes short term rigidities and the market power of

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[1] As a survey see Rogoff (1996).
[2] See, for example, Asea and Mendoza (1994) and De Gregorio and Wolf (1994).

importing companies. These market imperfections allow foreign suppliers to set the markup of prices over the marginal cost. The assumptions of pricing-to-market approach, however, are in many ways under much debate and not necessarily even sufficient for long-term deviations in the aggregate price index analyzed in this study. An aggregate price index, such as a consumer price index, definitely includes goods produced by the industrial sectors which are best characterized by incomplete international competition but also industrial sectors which are almost competitive. Thus, we must further extend our toolkit.

Among all other economic explanations for incomplete pass-through we have limited this paper to the (international) real option investment theory inspired by MacDonald and Siegel (1986), Pindyck (1988), Dixit (1989a and b) and Dixit (1993). Using a real option theory we examine foreign firms' entry and exit decisions in competitive domestic importing markets.

Since foreign firms in focus, exchange rate uncertainty is considered. At first, a nominal exchange rate is determined by net capital flows. The value of the nominal exchange rate is assumed to be revealed before the firms make their decision whether or not to be in the market. We explicitly consider the effects of profit maximizing foreign firms' entry and exit decisions on the domestic tradable prices through changes in supply.

Dixit (1989a and b) assumes that the exchange rate follows a geometric Brownian motion. The expected uncertainty is introduced using historical variance, i.e. it is assumed to be an almost constant variable over time. The inaction band (no entry or exit) around the base value is either stable or determined by the market share of foreign firms. Although the values given for entry and exit trigger points in Dixit (1989b) are realistic, especially in the case of the median firm, we argue that the dynamic structure of the pass-through process is at least partly problematic. Deviations from the parity level are too volatile and persistent to be explained solely by the changes in the number of firms or small changes in expected uncertainty if uncertainty is generated by the volatility of a continuous stochastic process. To avoid these problems we extend the discussion, also considering possible jumps in the expected exchange rate time path.

The estimates in Sollis *et al.* (2002) show that there is a weaker mean reversion when the US dollar real exchange rate is overvalued relative to historical averages. The adjustment to PPP level is not only nonlinear (inaction band around mean) but also asymmetric. This is probably due to large appreciation in the 1980's, which is widely considered to be a temporary speculative bubble. Thus,

we especially consider entry decisions after a large appreciation. Moreover, we assume that the large exchange rate appreciation is expected to be temporary and the adjustment towards the parity level is expected to be relatively rapid. This gives us an opportunity to analyze the possibility that the US exchange rate follows the mixed Brownian motion-Poisson jump process after a large appreciation. Together with the real option investment theory this seems to offer an interesting explanation for the time path of the US dollar real exchange rate. After a large positive shock in nominal exchange rate foreign suppliers do not completely adjust their supply since there is a substantial likelihood of a large negative shock. We are now able to explain large and persistent deviations from the parity level, which are not constant in magnitude without assuming market imperfections or a systematic link between trade flows and the deviations of the real exchange rate from the trend.

2. TRADABLE GOODS PRICES AND THE US DOLLAR REAL EXCHANGE RATE

Traditionally tradable prices are assumed to follow the rule of one price. In this section we will discuss the collapse of the PPP for tradable goods. We examine more closely the time path of the US dollar real exchange rate and its deviations from the parity condition. Some special characters of the US tradable prices are also considered.

2.1. INCOMPLETE PASS-THROUGH

As discussed in Obstfeld and Taylor (1997), even the studies most favorable to long-run PPP suggest an extremely slow decay rate for international price differentials. The estimated half-lives for PPP deviations for most countries and time periods are found to be of the order of four to five years.[3] These estimates appear to imply more sluggishness than can be attributed solely to nominal rigidities.[4]

Persistent deviations from PPP indicate incomplete pass-through of the exchange rate for prices. One explanation commonly evinced for the incomplete pass-through puzzle is international price

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[3] See, for example, Wei and Parsley (1995) or Frankel and Rose (1996). Panel estimations, such as Papell and Theodoridis (1998) and Oh (1996), generally find somewhat more rapid reversion with half-lives of the order of 2 to 2,5 years.

discrimination. Krugman (1987) labeled the phenomenon of exchange rate induced price discrimination in international markets "pricing-to-market" (hereafter PTM). According to the PTM approach international markets for manufacturing goods are sufficiently segmented so that producers or retailers can, at least over some horizon, tailor the prices they charge to the specific local demand conditions prevailing on different national markets. Thus, firms set different prices for their goods across segmented national markets to compete with firms on those markets. Segmentation between national markets depends on substitution and transaction cost effects.

Since we analyze aggregate price index (CPI), it is somewhat problematic to take the ability to price discriminate through the substitution effect to be absolute. As emphasized by Rogoff (1996) segmentation might be the case for some goods, such as automobiles, where differences in national regulatory standards combined with a need for warranty service allow firms great leeway to price discriminate across countries. There is, however, a substantial amount of tradable goods which are homogenous in different countries. The findings of Knetter (1993) showing that pricing to market seems to characterize even the most mundane goods, are likewise not in line with this substitutability assumption.

Another possible explanation for persistency in price differentials is that traded goods markets are not completely integrated. Obstfeld and Rogoff (2000b) emphasize the importance of trade costs, which effectively limit price competition, in resolving the problem of persistent price differences.
It is possible that trade frictions, such as transportation costs, allow tradables prices to differ within some range without inducing profitable arbitrage. Estimates of average transport costs across all tradable goods range between 6 and 10 per cent.[5] In addition, tariffs and non-tariff barriers may cause important frictions. Thus, transaction costs should provide some scope for deviations from the law of one price. Based on this assumption there should be one rapid convergence band when price differences exceed transaction costs, and one slow or non-convergence band when price differentials are relatively small. Michael *et al.* (1997), and Obstfeld and Taylor (1997), among others, have found evidence that large PPP deviations die out more rapidly than small ones.[6]

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[4] Cheung and Lai (2000) reports hump-shaped adjustment paths of the US real exchange rates, which are also incompatible with standard sticky price models.
[5] See the discussion in Rogoff (1996).
[6] Michael *et al.* (1997) use long time series spanning 1791-1992 and postwar time series for bilateral exchange US real exchange rates. Obstfeld and Taylor (1997) use data measured relatively to the US after 1980.

The market organization which provides market power for firms, albeit essential for PTM, is under much research and debate.[7] In Cournot oligopolistic market formulation with homogeneous goods, Dornbusch (1987) summarized the elasticity of the equilibrium price with respect to the exchange rate as

$$\varphi = (n*/N)(xw*/p)$$

The elasticity formula has two determinants: the relative number of foreign firms and the ratio of marginal cost of foreign firms (w*) in home currency (x) to price of foreign suppliers in domestic currency (p)[8] The first determinant of the elasticity formula simply reveals that higher share of imports increases the elasticity of tradables price level for exchange rate changes. The second determinant of the elasticity formula is a price discrimination determinant. It is essential for the pricing-to-market literature, which has focused on the issue of markup adjustment as a possible explanation for very slow response of tradable goods prices to exchange rate movements.

The recent literature supplements the PTM assumption with an extra assumption regarding local-currency pricing, i.e. prices are assumed to be sticky in the local currency of the buyer.[9] It is again important to note that we concentrate on consumer price indices, i.e. trade at the consumer level. Something important must be happening between the consumer level, where the medium-term effect of exchange rates on prices is virtually zero for many goods, and the wholesale level, where price effects tend to be less than proportional but also significantly greater than zero.[10] According to Obstfeld and Rogoff (2000a), local currency pricing is pervasive at the retail level, which should explain these findings.[11] Price contracts are not, however, generally thought to be very long-lived.[12] Bergin and Feenstra (2001) found that price contracts combined with PTM are able generate endogenous persistence beyond the exogenously imposed rigidity. Nevertheless, they are not able to reproduce the degree of persistency observed in the data. [13]

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[7] See the discussion in Cheung *et al.* (1999).
[8] As shown in Dornbusch (1987), qualitative results do not depend strictly on the chosen model. Dornbusch (1987) examines the Cournot oligopolistic model and a more Keynesian type of Dixit-Stiglitz model.
[9] See, for example, Devereux (1997).
[10] See the results in McCarthy (1999).
[11] Obstfeld and Rogoff (2000a) argue that if local-currency pricing was common practice in manufacturing then a country's manufacturing terms of trade improve if its currency depreciates. This is inconsistent with the data.
[12] Typically at most one to two years.
[13] Chari *et al.* (1998), using only the exogenous persistence of price contracts, find that sticky prices can help replicate persistence, but only if at least 3 years' prices contracts are assumed.

Finally, we again would like to stress the importance of trade cost for incomplete pass-through. As pointed out by Obstfeld and Taylor (1997), advances in the theory of investment under uncertainty imply that a band of no-arbitrage should be interpreted as a resulting not only from concrete shipping costs and trade barriers, but also from sunk costs of international arbitrage.[14] Thus, one interesting perspective, which partly combines above arguments, is provided by the models that emphasize dynamic supply effects. Dixit (1989a and b) show that when firms face sunk entry costs when breaking into foreign markets, the extent of pass-through will depend on the expected changes in the nominal exchange rate. Different pricing behavior on different markets now depends on the entry and exit decisions of competitive firms. Prerequisites for entry into foreign market are, for example, investments in a marketing and distribution network, which are especially important at the consumer level.

2.2. US TRADABLE CONSUMER PRICES

There are some special issues related to US tradable prices. Canzoneri *et al.* (1999), using panel data, argue that the failure of PPP to explain traded goods prices is especially important for the US dollar. It is certainly true that a non-tradable component is important in the determination of tradables prices. Engel's (1999) results make clear, however, that more than this must be going on, since prices for tradable goods do not seem to respond any faster to exchange rate movements than do the prices of non-traded goods. Engel has also pointed out that tradables prices appear to account for a large part of the movement of US real exchange rates independently of the chosen real exchange rate between the US and other high income countries.[15]

It is also interesting to view the relation between the nominal exchange rate and the price differential. We use the US dollar/German mark variables as an illustrative example. The large deviations of nominal exchange rate from the long-run price trend are shown in Figure 1, where c is the US dollar/German mark nominal exchange rate and *dif* is the difference of consumer price indices in the U.S. and Germany.

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[14] O'Connell and Wei (1997), provide a continuous time model on goods arbitrage that highlights the relative importance of proportional and fixed cost of transactions. Their evidence indicates that the fixed component is dominant.

[15] Engel (1999) uses aggregate price indices, but there are also similar findings even for highly disaggregated data. See Giovannini (1988) and Engel and Rogers (1996).
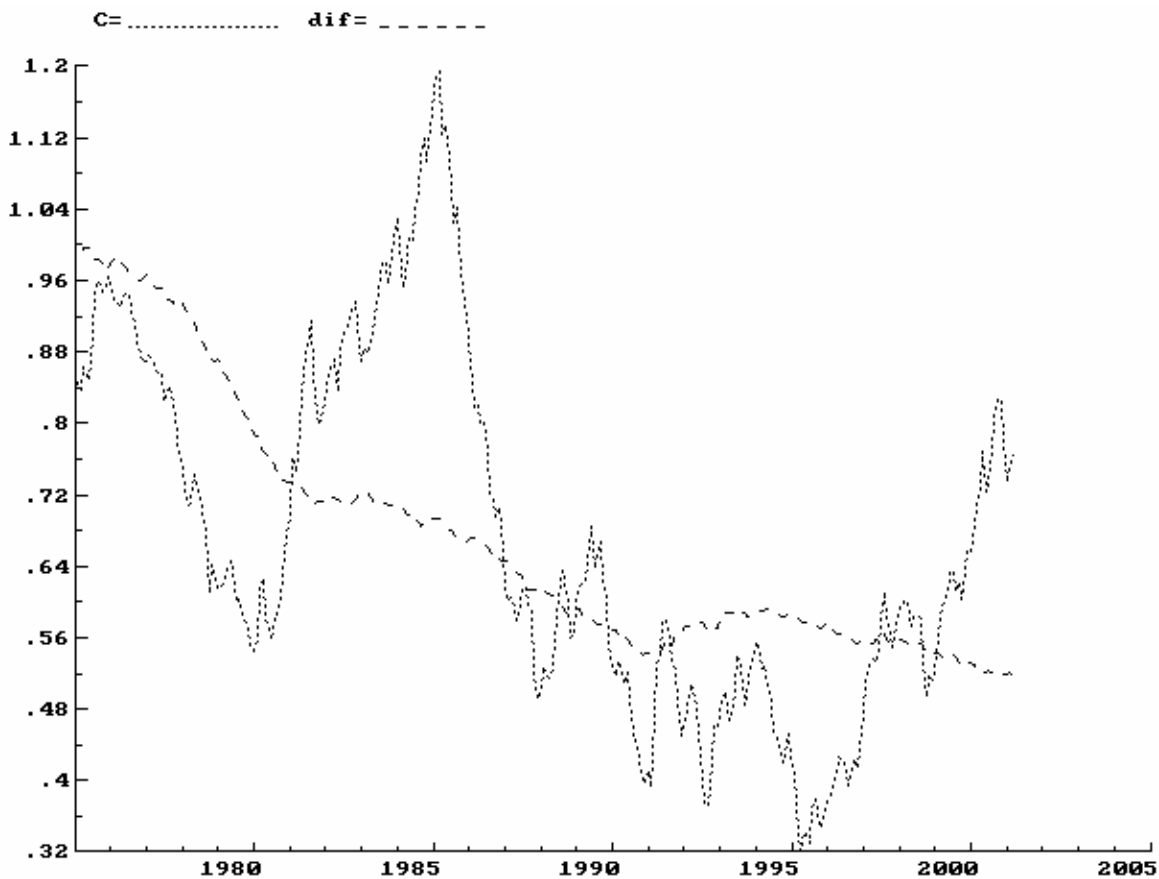
Figure 1. Nominal exchange rate and price differential.

Deviations from the parity level appear to be large and time varying. They are too volatile to be explained by the constant narrow convergence band, especially in the middle of the 1980s and probably also at the begin of the new millennium. There are also large jumps in the time path of the nominal exchange rate. Note that the large deviations also seem to be as persistent as nominal exchange rate swings.

Pass-through from the dollar exchange rate to US tradable prices fall in the 1980s.[16] The estimates in Sollis *et al.* (2002) also show that there is a weaker mean reversion when the US dollar real exchange rate is overvalued relative to historical averages. The puzzling fall of exchange rate pass-through during that time period of an appreciating dollar has been pointed out in various ways in the literature. Marston (1990) points out that nominal exchange rates surprises lead only to temporary changes in pass-through due to preset prices. Permanent changes are also possible if there are fundamental changes in PTM behavior. PTM behavior, in turn, is determined by the differences in

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[16] See, for example, Dornbusch (1987), Froot and Klemperer (1989).

price elasticities of demand curves in each importing country. Dixit (1989b) assumes that the sunk cost of entry decreases pass-through if exchange rate appreciation is not large enough. However, sufficiently large appreciation may raise pass-through permanently even if appreciation itself is only temporary.[17] Froot and Klemperer (1989) show that a model with consumer switching costs will lead exporters to respond differently to temporary and permanent changes in exchange rate. They examine the effects of temporary appreciation of the dollar focusing on dynamic demand side effects in an oligopolistic market. In their model temporary appreciation increases the foreign currency value of current, relative to future, dollar profits. When the dollar is temporarily high, foreign firms will find investments in market share less attractive, and will prefer instead to let their current profits increase.

2.3. US DOLLAR EXCHANGE RATE

As discussed in Goldberg and Knetter (1997), the assumption of temporary exchange rate change in Froot and Klemperer (1989) is somewhat problematic since the literature on exchange rate determination shows only very weak evidence in favor of reversion to PPP, meaning that most changes might be viewed as permanent. Thus, we take a different approach by assuming that the exchange rate follows a mixed jump diffusion process. Hence, we allow for the exchange rate to undergo unexpected, discrete changes. Especially after a large shock there is a certain likelihood of unexpected jumps in the process.

Obviously, we should first find some evidence for the mixed random walk process. The time period between 1980 and 1985 is particularly notable. As Froot and Klemperer (1989) note, nominal interest rate differential, a common measure of expected depreciation, shows that the dollar was expected to depreciate most rapidly in the early 1980s, just when the rate of appreciation was also the greatest. However, as pointed out by Juselius and MacDonald, (2000) the prolonged nature of the appreciation would seem to be unwarranted solely in terms of interest rate differential or the "safe haven" effect. They also argue that it might be a speculative bubble which was intensified by the role of the dollar as a reserve currency in the international monetary system.

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[17] Permanent effect can be cancelled only if the appreciation period is followed by a large depreciation period.

The above discussion provides an opportunity to study the order of pass-through after a large exchange rate appreciation. The appreciation is expected to be temporary and the adjustment towards parity is expected to be relatively rapid. Regardless of the source of the exchange rate shock the effects are similar if investors perceive even a remote likelihood of a large and rapid depreciation.

3. INTERNATIONAL INVESTMENTS AND REAL OPTION INVESTMENT THEORY

We restrict our analysis to foreign firms' investment decisions at the industry level. Foreign firms' entry and exit decisions are based on the expected discounted value of future profits of foreign investment. To determine the expected discounted value of future profits of foreign investment we use a real option investment theory originally developed by MacDonald and Siegel (1986), Pindyck (1988) and further developed by Dixit (1989a,b) and Dixit (1993). Real option investment theory is a solution method, which uses Itos's lemma for the analysis of a stochastic investment income process.

3.1. REAL OPTION THEORY

For entry into the market, the crucial importance in the real option theory is the moment in time when firms decide to invest in a single project. Firms own an option to enter the market at any moment in time. This option has an exercise price which is also a sunk cost of entering the market. The value from exercising the option is the expected present discounted value of future profits from serving that market. Since the value of investment is unknown, there is an opportunity cost to invest today. In terms of option theory, the investment rule can be stated as follows: invest when the value of the project exceeds its costs by an amount at least equal to the option value of waiting to invest.[18] Similarly, if firms consider to leaving the market, investments are typically at least partly irreversible, again incurring sunk cost to the firms.

In the real option literature it has been generally assumed that higher level of uncertainty increases option value and this leads to more distant critical value for option exercise, i.e. there is an inverse relationship between uncertainty and investment, since greater uncertainty increases the option

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[18] As discussed in Dixit and Pindyck (1994), this decision rule gives completely different results than a traditional net present value decision rule.

value of waiting.[19] Thus, fewer events of entry will be observed. This argument implicitly includes the assumption of concavity of the investment value, since we know from the general theory of choice under uncertainty that greater uncertainty will increase the expected value of an action if the payoff is convex in random variable, and decrease it if the payoff is concave.[20]

We examine the effect of uncertainty on foreign investment in competitive markets. The increasing volatility of a geometric Brownian motion means, however, that higher price levels may be achieved. Caballero (1993) noted, that under perfect competition between firms where the price elasticity of demand is almost infinite, an increase in existing prices raises the value of the investment both directly through the price change and also indirectly through the increase in optimal output. The latter effect would induce greater amount of investments, i.e. lead to the positive relationship between uncertainty and investments.[21] However, Dixit (1993) argued that even in the case of perfect competition the level of price about which to make the demand more elastic is itself endogenous and, moreover, acts as a ceiling or reflecting barrier. More precisely, a competitive firm's investment decisions are restricted by the price ceiling at a certain price level imposed by firm rational expectations of other firms' entry decisions. Thus, the ceiling barrier will make expected profits on investment concave in the price process and increase the option value of waiting if uncertainty is increased even when perfect competition is assumed.

3.2 A MODEL FOR INCOMPLETE PASS-THROUGH

To compare the effect of mixed process assumption on pass-through to the Brownian motion assumption in Dixit (1989b), we introduce a simple market structure following Dixit (1989b). In this model demand is assumed to be stable, i.e. the prices of imported goods are supply side determined. The supply side of economy is competitive, i.e. importing firms act as price-takers. Let the demand for import goods be determined by the inverse demand function $p = P(n)$. For all n there is a function such that $p = P(n) = U'(n)$, i.e. the n*th* firm's marginal contribution to utility is

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[19] See, for example Pindyck (1988).
[20] Although this definition is built on Jensen inequality and risk averse investor, it is possible to show that, under certain conditions, this is also true even for a risk neutral investor. See Dixit (1993).
[21] Note that Caballero's result is based on a model with convex adjustment costs. In addition, Caballero's result treats the firm in isolation.

equal to the market price P(n). Since U(n) is arbitrary to within an additive constant, we can set U(0) = 0. We can write

$$P(n) = U(n) - U(n-1) \qquad \text{and} \qquad U(n) = \sum_{j=1}^{n} p_j \qquad (3.1)$$

Only trade is considered, not the production location. Similarly, only the supply of importing firms with profits measured in foreign currency is considered, not the supply changes of domestic firms. It is also assumed that each foreign firm sells one unit of output per unit of time, i.e. supply changes depend strictly on changes in number of firms. Thus, price level is completely determined by the number of active foreign firms.

Foreign firms are characterized as follows. There are the same sunk costs for every foreign firm to enter the domestic market. Firms are allowed to differ in their variable costs.[22] Following similar assumptions as in (3.1), the variable costs of the first n firms are measured as

$$w_n = W(n) - W(n-1) \qquad \text{and} \quad W(n) = \sum_{j=1}^{n} w_j \qquad (3.2)$$

The firms are labeled so that w_n is increasing. Since w_n is increasing, W(n) is increasing and convex. The variable costs of foreign firms are measured in foreign currency. All foreign firms also have same relative technological progress, i.e. they differ only in their variable costs.
Foreign firms are risk neutral and have rational expectations, i.e. they maximize the expected present value of profits in foreign currency. The maximand is

$$E\left\{ \int_0^\infty \{X_t U(n_t) - W(n_t)\} e^{-\rho t} dt - \sum_i X_i \{k[\Delta n_i]\} e^{-\rho i} \right\} \qquad (3.3)$$

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[22] Sarantis (1999) has argued that the transition between the convergence regimes based on total transaction costs is smooth. This is probably due to heterogeneous investors. Thus, in this section we introduce an international real option investment theory, which stresses that heterogeneous investors face sunk costs of entry when breaking into foreign market.

where X is nominal exchange rate and ρ is the real interest rate, which is used as the discount rate by the foreign firms. At instant t = i, when the numbers of foreign firms change, there is a sunk cost k for every foreign firm to enter the market.

Nominal exchange rate X_t determines the price level through the investment decision of foreign importing firms. At the beginning of the period the size of net capital inflows is revealed. In order to assess the simple feedback from net capital inflows to the nominal exchange rate, we assume that exchange rate is completely determined by net capitals inflows in the medium run. Thus, there is no feedback from firms' investment decisions to the nominal exchange rate process, i.e. the exchange rate process is exogenous in respect of the investment decisions of foreign exporting firms.

After the size of net capital inflows has been revealed firms decide whether to enter or exit. Depreciation of foreign currency increases the demand for foreign goods. There are also necessary sunk costs for entry in home markets. Thus, investment decisions (export decisions) are based on the maximand given above. In a time interval when no entry or exit takes place, supply is fixed and prices are proportional to the nominal exchange rate shock. Maximization of the expected present value of investments, in turn, determines a range of current values of the exchange rate that will lead no firms to either enter or exit. Thus, prices are determined due to the competitive risk neutral foreign firms' entry and exit decisions.[23]

3.3. THE MIXED POISSON JUMP AND GEOMETRIC BROWNIAN MOTION PROCESS

The time path of the exchange rate follows a random walk process, but we also allow for the possibility that, at some random point in time, the time path will take a Poisson jump. In the following analysis we use the appropriate version of Ito's Lemma, which combines a random walk and Poisson jump effect. We discuss the expected time series process after a large appreciation.

The random walk and Poisson jump effect is, more accurately in a continuous time representation, a mixed geometric Brownian motion and Poisson jump process, where the former is continuous and the latter occurs infrequently. We denote the process as

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[23] We underestimate the pass-through since we do not consider the expansion of firms already established in the market.

$$dX = \alpha X dt + \sigma X dz + X dq$$

where X is a non-stationary exchange rate variable at time t, σ is the standard deviation of the process and dX is the infinitesimal change in X over the infinitesimal time interval dt. Without any jumps in the process the stochastic change in the variable over this interval of time is dz. The normally distributed random variable dz represents exogenous random shocks to the change in X over a small time interval dt. The expected value of dz is zero. dq is the increment of a Poisson process. dq and dz are independent so that E(dz*dq) = 0. The conditionally expected value of dX is equal to the sum of the deterministic component, $\alpha X dt$ and the conditionally expected value of the Poisson component $XE(dq)$. We assume that if an event occurs, q changes by some fixed percentage θ. The increment of a Poisson process takes the form such that:

$$dq = \begin{cases} 0 & \text{with probability}: (1-\lambda)dt \\ \theta & \text{with probability}: \quad \pi\lambda dt \\ -\theta & \text{with probability}:(1-\pi)\lambda dt \end{cases}$$

where the probability of the Poisson jump effect is λ and π is the probability of a positive jump.

The decision to invest is equivalent to deciding when to exercise an option. Thus, we can analyze a firm's investment behavior by analyzing the option value of investment $H_n(X)$. We provide a full derivation of the problem in Appendix 1. It closely follows the cited literature. However, two important modifications have been made. First, the time path of the expected value of the investment now follows a mixed process, since the exchange rate may be affected by random Poisson jump shocks. Second, the market determined expected rate of return is assumed to remain constant, i.e. the higher probability of a sudden fall of the investment income is accompanied by an increase in α.

Ito's Lemma gives us the following differential equation

$$\begin{aligned} &\tfrac{1}{2}\sigma^2 X^2 H_n''(X) - rH_n(X) + (r - \pi\lambda\theta + (1-\pi)\lambda\theta - \delta)XH_n'(X) + XU(n) - W(n) \\ &+ \pi\lambda H_n(X) - (1-\pi)\lambda H_n(X) - \pi\lambda H_n[(1-\theta)X] + (1-\pi)\lambda H_n[(1-\theta)X] = 0 \end{aligned}$$
(3.4)

where U(n)-W(n) is the flow of future dividends, r is a real interest rate, and δ is the difference between α and r.

We are especially interested in the entry decision after a large shock in exchange rate value. Thus, we only consider entry decisions of foreign firms and exclude exit decisions. Since we assume that the firm is risk neutral, its discount rate is equal to the interest rate, ρ=r. Assuming firms never shutting down, Equation (3.4) has a general solution

$$H_n(X) = B(n)X^\beta + (XU(n)/(r - \alpha - (\pi\lambda\theta - (1-\pi)\lambda\theta))) - W(n)/r.$$ (3.5)

The last two terms on the right hand side of (3.5) give the expected present discount value of maintaining exactly n firms forever, starting with the exchange rate X.[24] Then, the first term must be the value of the options to change the number of firms. B(n) is constant to be determined and β is a root of the quadratic equation.

$$f(\beta) = \tfrac{1}{2}\sigma^2\beta(\beta-1) + (r - \pi\lambda\theta + (1-\pi)\lambda\theta - \delta)\beta + (\lambda\pi - r - \lambda(1-\pi))$$
$$- \pi\lambda(1-\theta)^\beta + (1-\pi)\lambda(1-\theta)^\beta = 0$$ (3.6)

For sensible parameter values there is one root to Equation (3.6), which is greater than unity. This root together with constant B(n) gives the exchange rate values for the entry of foreign firms.

As shown in Dixit (1989b), if I_n is the optimal exchange rate for entry then the relationship between the investment value at which it is optimal for the *n*th firm to invest and the sunk cost of the investment is given by the value matching condition

$$H_{n-1}(I_n) = H_n(I_n) - kI_n,$$ (3.7)

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[24] Note that $E(X_t | X_0) = X_0 \exp(\alpha t + (\pi\lambda\theta - (1-\pi)\lambda\theta)q$, where X_0 is an initial exchange rate. Therefore

$E\int_0^\infty \{X_t U(n) - W(n)\}\exp(-rt)dt = \dfrac{XU(n)}{r - \alpha - (\pi\lambda\theta - (1-\pi)\lambda\theta)} - \dfrac{W(n)}{r}$. See the discussion in Dixit (1989b).

where k is an entry cost. This endogenizes the value of the investment and the number of firms. The smooth pasting condition is

$$H'_{n-1}(I_n) = H'_n(I_n) - k \qquad (3.8)$$

Substitute the functional form of the solution from (3.5) into (3.7) and (3.8). Define $b_n = B(n-1) - B(n)$ for n=1,2,…, N firms. We obtain

$$-b_n I_n^\beta + (I_n p_n / (r - \alpha - (\pi \lambda \theta - (1 - \pi) \lambda \theta))) - w/r - kI_n = 0 \qquad (3.9)$$

and

$$-\beta b_n I_n^{\beta-1} + (p_n / (r - \alpha - (\pi \lambda \theta - (1 - \pi) \lambda \theta))) - k = 0 \qquad (3.10)$$

There are now two equations for two unknowns; I_n and b_n.

4. ECONOMIC INTERPRETATION OF REAL OPTION THEORY

Above we have argued, using a real option investment theory that sunk costs of investment and uncertainty affect investment decisions. In this section we first discuss micro data evidence for this relationship and then examine the importance of the nature of risk for investment behavior.

4.1. FOREIGN IMPORTING FIRMS AND EXCHANGE RATE RISK

Empirical research on the importance of uncertainty and sunk cost hysteresis in real exchange determination has typically focused on asymmetries in the response of import and export prices to exchange rate changes without explicitly focusing on the entry and exit decision into the market. Baldwin (1988), for example, examines the reactions of the US import prices to exchange rate changes. He finds a structural shift in the relationship between US aggregate import prices and the dollar appreciation during the early 1980s. Wei and Parsley (1995) cannot find exchange rate volatility to be a significant predictor of trade flows between the U.S. and Canada. Campa and

Goldberg (2002) also find that exchange rate volatility has no role in explaining exchange rate pass-through changes.[25]

Indeed, a large amount of theoretical work on firms' entry and exit decisions contrasts with the very slight empirical evidence for the behavior of foreign firms.[26] Since explicit entry decisions into the U.S. market are important for our analysis, an interesting study is provided by Campa (1993). He focuses on the effect of exchange rate changes on decision to enter the US market. The empirical estimates confirm an inverse relation between the exchange rate uncertainty and entry decisions. However, Campa (1993) was not able to discover a relationship between exchange rate trend (α) and the number of firms deciding to enter the market.

4.2. RISK AND FOREIGN FIRMS

As discussed in section 2, deviations of the exchange rate from its parity level seem to be time varying in magnitude. In Dixit's (1989b) model with geometric Brownian motion, this can be explained by either large and persistent changes in expected risk or large and persistent changes in relative number of foreign firms. The time path of the expected risk defined as historical variance is illustrated in Figure 2, which plots quarterly series of the differences of the US dollar/German mark real exchange rate. The importance of changes in the relative number of foreign firms is, in turn, shown in Figure 3.

In general, the quarterly change of the (log of the) exchange rate is a stable variable for most of the time.[27] Although the value 0,04 is a relatively rare event, there are a few large peaks which in order of magnitude are almost 0,06. Thus, the time path of the exchange rate is not best characterized by the high and low level regimes of uncertainty but by the few large peaks in the time series. Although there is no unique way in which firms form expectations on risk, it might be reasonable to assume that the overall expected risk combines the effects of the continuous motion parameter σ and the jump parameter dq.[28]

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[25] Campa and Goldberg (2002) find out that pass-through into import prices is lower for countries with low exchange rate variability.

[26] Roberts and Tybout (1997) analyze exporting experience on the decision of Columbian manufacturing plants to participate in foreign markets and find support for the importance of sunk costs. Campa (2000) looks at the response of a country's export supply to exchange rate changes using Spanish data. Findings in Campa (2000) support the importance of sunk cost hysteresis in entry behavior but find exchange rate uncertainty unrelated to export supply.

[27] The GARCH(1,1) model, for example, is not statistically acceptable.

[28] Common assumptions are perfect foresight or static expectations based on historical data.
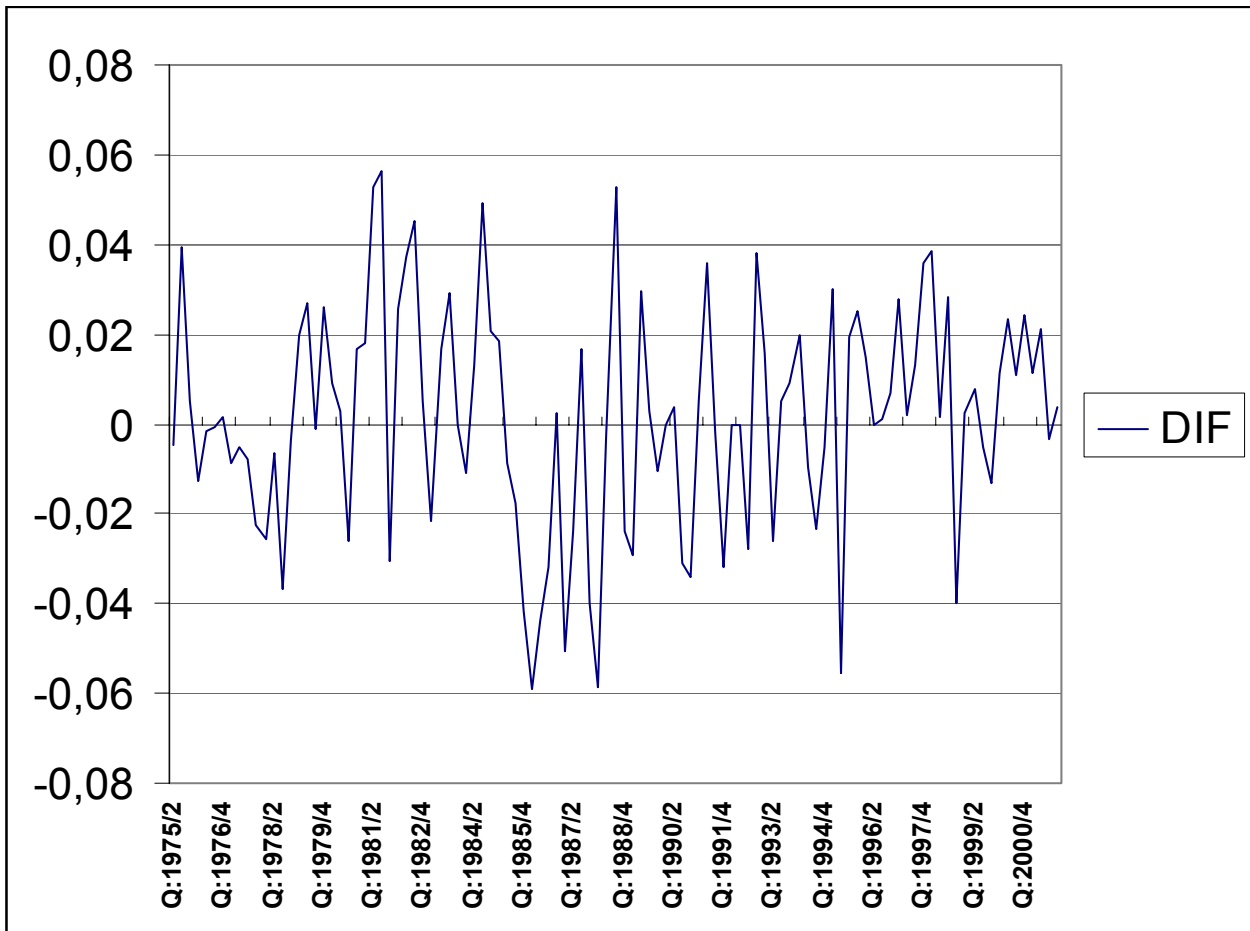
Figure 2. Quarterly change of the real exchange rate.

The effect of large and persistent changes in relative number of foreign firms can be seen in Figure 3 below. As an expression for this effect we have used the ratio of imported goods to private consumption of goods (s). n is again the US dollar/German mark real exchange rate. The import share has been increasing quite steadily during the last two decades. This is probably due to the structural changes on global goods market.[29] The increased share of imports should increase the pass-through as discussed in section 2. However, the changes in imports are not highly correlated with changes in magnitudes of deviations from the fundamental PPP level.[30] During some subperiods with an appreciating dollar, the import share of private consumption goods increases more than just by trend but during the other subperiods the effect of appreciation on imports shares is relatively small.[31]

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[29] See the discussion in Baldwin and Krugman (1989).
[30] See also Figure 1.
[31] Campa and Wolf (1997) also find (using G7 data) that the deviations of import share from trend do not systematically co-vary with the deviations of real exchange rates from trend.

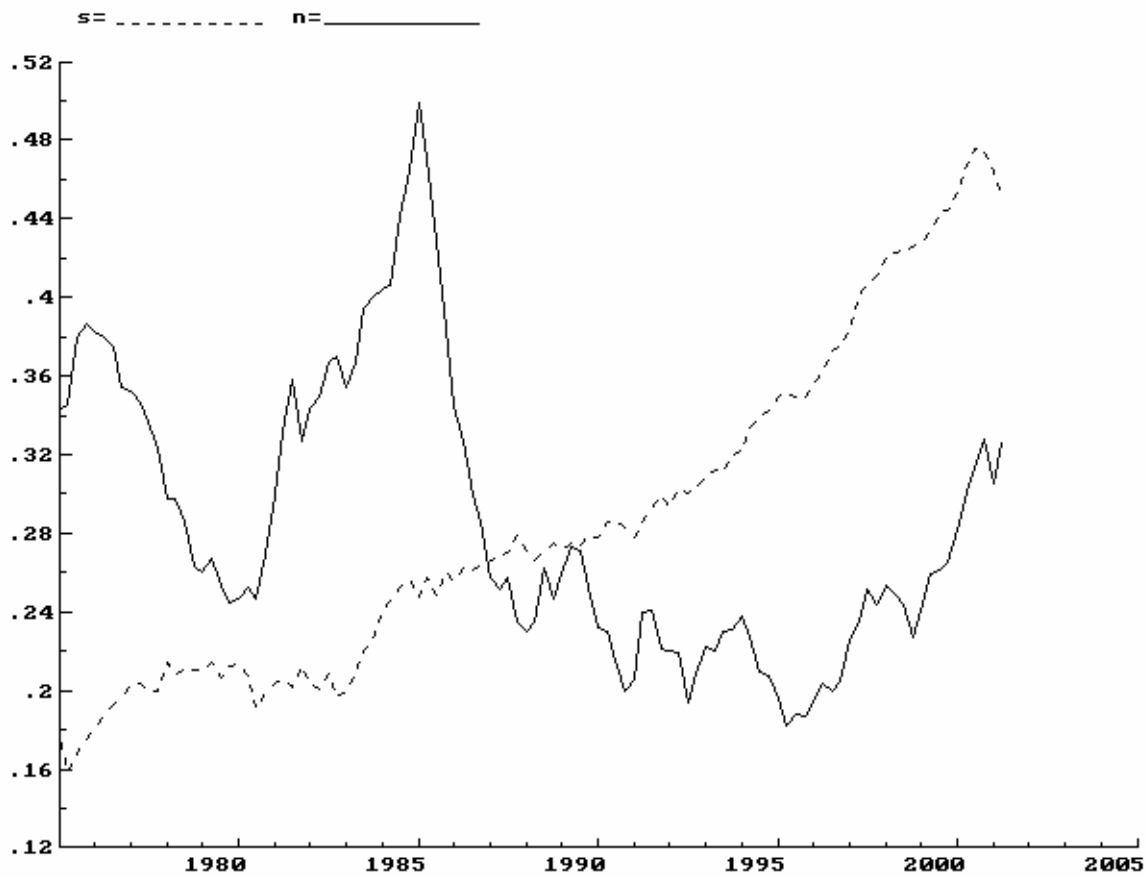s= _ _ _ _ _ _ _ _   n=_____

 Figure 3. Import share.

In the following analysis we will examine numerically the effects of the mixed process on the time path of the real exchange rate. This may help us to understand deviations from the parity level which are not an order of magnitude constant.

4.3. CALCULATIONS

An interesting economic interpretation of the above equations is given by examining the sensitivity of the results on parameters changes. We calculated the dependence of I_n on θ and λ. To compare our results with Dixit (1989b) we make similar assumptions on firms and use similar domestic demand function for imports. (For details see Appendix 2). Since we do not allow for exit we increase the entry values. Our primary aim, however, is to examine the effect of the time series behavior assumption on pass-through, not how wide a band of inaction is *per se*.

At first we assume that the annual standard deviation is 0,10, which is approximately a standard deviation of the US dollar real exchange rate. The probability of negative jump is assumed to be very high (1-π=0,8) since we examine investment behavior after a large and temporary shock. The interest rate is two and half per cent, the sunk cost of investment is 2 and the maximum number of foreign firms is n = 100.[32] We calculate the results for the median firm (n = 50). The expected time until the next jump is 1/λ.

Table 4.1. Size and time. Stdv. 0,10.

	time	0,125	0,25	0,5
size	0,1	1,58	1,6	1,66
	0,2	1,65	1,74	1,91
	0,3	1,76	1,91	2,33

Table 4.1 presents the results. The critical value is higher the shorter the expected time to the next shock. The effect is more pronounced when the size of the shock increases. If the Poisson jump occurs frequently, it is reasonable to assume that foreign importers will wait it out. If a geometric Brownian motion is assumed, the critical value of investment needs to be 1,55, i.e. 55 per cent above the stationary value 1.[33] The critical value is very close to a geometric Brownian motion value if the parameters λ and θ are assumed to be at the minimum.

Above we have assumed that heterogeneous investors face sunk entry cost k=2 when breaking the foreign markets. The annualized sunk cost is then five percent of the full cost for a median firm. Since the amount of sunk costs for entry into the foreign market at the consumer level may vary significantly, we have also made calculations based on the assumption k=1, i.e. the annualized sunk cost is two and a half percent of the full cost for a median firm. This assumption decreases I_n from 1,74 to 1,70 if the size of expected shock is 0,2 and the time 4 years. A geometric Brownian motion assumption now gives the value 1,50 for entry. Thus the results are not very sensitive on this assumption.

The sensitivity of the results can also be discussed also in the light of the assumptions made about risk. In the following analysis we have decreased standard deviation from 0,10 to 0,05. The results are reported in Table 4.2 below.

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[32] We also assume that α=0.
[33] If we allow exit the critical value is 1,25. The net present value (NPV) decision rule gives values very close to one.

Table 4.2. Size and time. Stdv 0,05.

	time	0,125	0,25	0,5
size	0,1	1,28	1,32	1,38
	0,2	1,37	1,49	1,69
	0,3	1,50	1,73	2,12

For a lower level of standard deviation, the critical values also increase as parameters θ and $1/\lambda$ increase. The combined effect of size and time is now larger than with a higher continuous risk assumption. If a geometric Brownian motion is assumed, the entry value is 1,26.

As the standard deviation of the real exchange rate is 0,1 with the geometric Brownian motion assumption, we obtain the value 1,55, which is approximately the same value as that obtains by assuming the standard deviation 0,05 and a thirty percent expected drop during the next eight years or a twenty percent drop during the next four years as presented in Table 4.2. Thus we obtain a similar effect by assuming higher expectations for standard deviations as by assuming a mixed process.

4.4. EXCHANGE RATE PASS-THROUGH

First, we examine an exchange rate pass-through assuming the exchange rate lies on stationary value 1 with the standard deviation 0,05 and the median case, market share 25%, $p = 1$ and $w = 0,95$. The size of the shock is 0,2 and the expected time four years. As long as the number of foreign firms is constant at n, there is no pass-through at all. This is 49% above the stationary value assuming a mixed process and 26% assuming geometric Brownian motion. When the exchange rate appreciates even more 10 per cent from 1,49 to 1,64 prices change six percent and the pass-through is 0,6. We also obtain very similar results if the number of foreign firms is 70, i.e. market share is 32%. Then the exchange pass-through is 0,55.

5. CONCLUSIONS

Our quantitative results are far from perfect. We did not allow exit. This increases the critical value needed for entry. We have also assumed that foreign firms sell one unit of output per unit of time.

In reality, however, firms are not tied to selling just a unit output as we have assumed. More flexible operational space would decrease the exchange rate level, which is necessary for the positive expected income from investment. By assuming unit output we also underestimate an expansion of the previously established firms. As discussed in Dixit (1989b), this is an important phenomenon especially if the exchange rate takes values high enough for entry. Thus, the exchange rate pass-through is probably close to one in the phase with entry.

The most important contribution is not, however, the precise numerical result but the general dynamics of the adjustment process to the parity level. We can conclude that the necessary level of exchange rate for profitable investment may depend on expected Poisson shock and this, in turn, allows long run deviations from the purchasing power parity level if the source of the expected Poisson shock is persistent. We are now able to explain large and persistent deviations from the parity level, which are not constant in magnitude without assuming market imperfections or a systematic link between trade flows and the deviations of the real exchange rate from the trend.

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APPENDIX 1

The outcome of the maximization problem (3.3) can be written as $H_n(X_t)$. Since we assume that the firm is risk neutral, its discount rate is equal to the interest rate, ρ=r. We can write

$$H_n(X_t) = E\left\{\int_t [X_s U(n) - W(n)]e^{-r(s-t)}ds - \sum_i X_i[k\Delta n_i]e^{-ri}\right\} \tag{1}$$

Over the infinitesimal time interval, $\Delta t \to 0$, the Bellman equation for $H_n(X_t)$ is

$$rH(X_t)dt = [X_t U(n) - W(n)]dt + E_t[dH_n(X_t)] \tag{2}$$

The sum of the flow dividend term and the expected capital gain is equal to the expected normal return. As discussed in section 3, we assume that the expected value of the foreign investment follows a mixed process after a large appreciation. The stochastic process of mixed geometric Brownian motion and Poisson process is described by the equation

$$dX = \alpha X dt + \sigma X dz + X dq \tag{3}$$

where the Poisson process (dq) takes the form given in section 3.

Over an interval of time where the number of foreign firms remains unchanged, the evolution of $H_n(X_t)$ is given by Ito's lemma.

$$dH_n(X_t) = \left[\left[\alpha X H_n'(X_t) + \tfrac{1}{2}\sigma^2 X^2 H_n''(X_t)\right]dt + \sigma X H_n'(X_t)dz \right.$$
$$\left. + \left\{\pi\lambda[H_n(X_t) - H_n\{(1-\theta)X_t\}] - (1-\pi)\lambda[H_n(X_t) - H_n\{(1-\theta)X_t\}]\right\}dt \tag{4}$$

Since the variance term of the Brownian motion (dz) is defined as $dz = \in \sqrt{dt}$, where $\in$ is a normally distributed random variable with a mean of zero, the expected value of dz is zero. Thus, taking expectations we can write

$$E[dH_n(X)] = \alpha X H_n{}'(X)dt + \tfrac{1}{2}\sigma^2 X^2 H_n{}''(X)dt +$$
$$\{\pi\lambda[H_n(X) - H_n\{(1-\theta)X\}] - (1-\pi)\lambda[H_n(X) - H_n\{(1-\theta)X\}]\}dt \qquad (5)$$

α should be less than the discount rate of risk neutral firm, r, otherwise it is always optimal rather to wait than to invest. Let δ denote the difference r-α, δ>0. As pointed out by Dixit and Pindyck (1994, p. 172), we may assume that the market determined expected rate of return should remain constant (relative to the Brownian motion case) although there is now a certain possibility of negative shock. Then α-λ should remain constant because higher probability of a sudden fall of the investment income is now accompanied by an increase in α. Replacing α with $r - \pi\lambda\theta + (1-\pi)\lambda\theta - \delta$ and rearranging terms, then the above equation can be rewritten as

$$\tfrac{1}{2}\sigma^2 X^2 H_n{}''(X) - rH_n(X) + (r - \pi\lambda\theta + (1-\pi)\lambda\theta - \delta)XH_n{}'(X) + XU(n) -$$
$$W(n) + \pi\lambda H_n(X) - (1-\pi)\lambda H_n(X) - \pi\lambda H_n[(1-\theta)X] + (1-\pi)\lambda H_n[(1-\theta)X] = 0 \qquad (6)$$

In this case the increased probability of negative shock would be equivalent to an increase in the risk free interest rate.

The solution for the values of options to change the numbers of a firm takes the form $B(n)X^{\beta}$. The derivatives are

$$H_n{}'(X) = \beta B(n)X^{\beta-1} \qquad \text{and} \qquad H_n{}''(X) = \beta(\beta-1)B(n)X^{\beta-2}$$

Inserting above derivatives into Equation (6), we obtain a quadratic equation:

$$\tfrac{1}{2}\sigma^2\beta(\beta-1) + (r - \pi\lambda\theta + (1-\pi)\lambda\theta - \delta)\beta + (\pi\lambda - r - (1-\pi)\lambda)$$
$$- \pi\lambda(1-\theta)^{\beta} + (1-\pi)\lambda(1-\theta)^{\beta} = 0 \qquad (7)$$

APPENDIX 2

The maximum number of foreign firms is N=100. The net import demand function is

$$Q = 250 - 200p$$

which gives the price equation

$$p_n = 1{,}25 - n/200$$

and market share

$$s_n = n/(160 + 0{,}8n)$$

The profile of variable cost is increasing and takes the form

$$w_n = 0{,}85 + n/500$$

For the median firm

$$w + \rho k = 1 = p$$

PURCHASING POWER PARITY PUZZLE:
A SUDDEN NONLINEAR PERSPECTIVE

ABSTRACT

The aim of this paper is to construct a simple nonlinear model for the U.S. dollar – euro real exchange rate. The nonlinear model considered allows the adjustment towards long-run equilibrium to be sudden as well as smooth. We found that the adjustment is sudden.

1. INTRODUCTION

We examine the U.S. dollar - German mark (euro) real exchange rate over the period 1982 to 2003. The bilateral exchange rate between the United States and Germany was chosen in order to fully understand the complicated interrelationship between USA and Europe. Because of the different economic structure of the euro area relative to Germany and also the different operational objectives of the European Central Bank relative to the Bundesbank, the euro may, of course, have different properties compared to the German mark. Nevertheless, there is good reason to believe that the study of the dollar-mark system is useful in terms of indicating how the exchange rate between two large economies may be determined in the market.

The keynote in our model is that the deviations from the purchasing power parity (PPP) equilibrium are stationary. Although the stationarity assumption is theoretically plausible between two industrialized countries, it is a common finding in an international finance literature that real exchange rates contain a unit root for the recent floating period. If the real exchange rate is, after all, a non-stationary variable, standard inference is not valid and our findings can be questioned. Thus, we should consider a stationarity assumption rather more carefully.

While we build on the stationarity assumption, we also allow for the possibility that the real exchange rate follows non-linear time series process. A number of studies have utilized variants of threshold autoregressive models to capture the nonlinear behavior of the real exchange rate. Michael et al. (1997), Baum et al. (2001), Taylor et al. (2001) and Sollis et al. (2002) apply smooth transition autoregression (STAR) models. They provide considerable evidence of nonlinear real exchange rates. Intuitively, nonlinearities in the real exchange rate imply that frictions in international trade result in bands within which relative international prices can fluctuate without a strong tendency to adjust towards the parity level.

PPP, however, is somewhat distinct from the pure law of one price concept applied to commodities. Our view is that PPP can be examined more properly in the domain of monetary economics and macroeconomic theory of inflation. What we attempt to do in this paper is to show that the central bank has an important role in an exchange rate system when the real exchange rate is measured by

the CPI deflator. The central bank is assumed to respond to deviations of the expected inflation and output from their desired levels, i.e. the central bank follows a Taylor rule.[1]

Kilian and Taylor (2003) state that rational market agents take stronger positions against exchange rate levels far away from the latent PPP equilibrium. Based on the assumption of heterogeneous agents, they consider smooth transition models. Since the role of the central bank may be crucial, it is desirable to allow the adjustment towards long-run equilibrium to be discrete as well as smooth. We use a nonmonotonic second-order logistic smooth transition autoregressive model (LSTAR2) to investigate nonlinear mean reversion. This a special case of a three-regime switching regression in which the two outer regimes are equal.

Local unit root behavior is consistent with the findings that the real exchange rate series fluctuate without a strong tendency to return to the equilibrium level when they wander in the neighborhood of it. We find using a second-order STAR that most of the observations belong to this regime. Large deviations from the parity level return relatively quickly towards parity level. This may lead to somewhat different conclusions concerning nonstationarity. The results of our analysis would suggest that when allowing for a non-linear alternative the real exchange rate may follow a globally stationary process and standard inference should hold.

We further found that the adjustment is sudden. This is probably due to the central bank's operations. Furthermore, there is a relatively large reference band around the parity level. This implies that the weight put on the U.S dollar-German mark (euro) real exchange rate in the monetary policy rules is low.

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[1] Several papers have added terms in exchange rates to otherwise standard Taylor rules. See, for example, Clarida et al. (1998), Engel and West (2004).

2. THEORY

The central bank is assumed to respond to deviations of expected inflation and output from their desired levels, i.e. monetary policy rules in the foreign and home countries follow a Taylor rule. Consistent with a Taylor rule, short-run interest rate will respond to the deviations of the nominal exchange rate from the underlying PPP target. The weight put on the real exchange rate in the monetary policy rule is positively related to the pass-through from the exchange rate to import prices. Frictions in international trade imply that the weight put on the real exchange rate in the monetary policy rule may be lower close to PPP than far away from it.

In Kilian and Taylor (2003) a consensus is gradually built among market agents that the exchange rate is misaligned as the exchange rate moves away from the equilibrium. This ensures a non-linear mean reversion of the exchange rate towards the PPP equilibrium. The expected variance of the real exchange rate is restricted by heterogeneous market agents' expectations and the real exchange rate time series follows a random walk process only in the neighborhood of PPP.

Even though individual market agents may firmly believe the exchange rate to be misaligned, they hardly have the power or resources to break the trend. If the exchange rate persistently disconnects from its fundamental PPP value, market agents may also lose their confidence in the usefulness of the PPP trading rule.[2] Then unanticipated monetary policy shocks may be necessary to break a trend. Sarno and Taylor (2001) suggest that even sterilized interventions may be useful once the exchange rate has moved a long way away from the fundamental equilibrium. Publicly announced interventions can now be seen as fulfilling a coordinated role in that they encourage individual market agents to enter the market at the same time.

It is reasonable to assume that the central bank is expected to extend only to monetary shocks that do not threaten its primary objective of low deviations of domestic target variables from their desired levels. Thus, publicly announced interventions signal future monetary policy intentions only if the exchange rate is relatively far away from parity level. If this distance from parity level is stable, market agents may recognize it. This implies a homogeneous response to large deviations from the PPP equilibrium.

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[2] Cheung and Chinn (2001) found that at the six month horizon 81 % of traders view PPP as irrelevant. At the very long horizon only 40% of traders agree that PPP has some influence.

3. EMPIRICAL NONLINEAR REAL EXCHANGE RATE MODEL

We do not primarily attempt to identify the effects of shocks. Empirically we do, however, aim to carefully elaborate the idea that the low weight put on the real exchange rate results in a band within which the real exchange rate time series follows a random walk process. Exchange rate levels far away from the PPP equilibrium are less likely to be persistent. Thus, the real exchange rate is a stationary variable which follows a nonlinear process. Uncertainty regarding expectations implies that a transition between two regimes may be discrete as well as smooth.

Most of the time the real exchange rate fluctuates within a band around the PPP equilibrium. The size of the band is determined by market agents' expectations of future monetary shocks. Inside the band the exchange rate follows a random walk and the best traders can do is to forecast no change. The fluctuation around the parity level may be interrupted by rare periods of turbulence, when the real exchange rate, under the influence of a succession of random shocks in the right direction, crosses over to the other implicit boundary. At such times, the deviations are expected to be temporary and to adjust towards parity.

3.1 THRESHOLD METHODOLOGY

When using a nonlinear methodology for analyzing price convergence, a discrete threshold methodology is found to be appropriate in one good world. Obstfeld and Taylor (1997), for example, use a self-exciting autoregressive (SETAR) model where the reversion is towards the edge of the band. They identify reasonable convergence speeds for disagregated tradable goods baskets.[3]

Outside this simple analytical structure, the specification of fixed thresholds becomes problematic. Moreover, when the real exchange rate is measured using price indices made up of goods prices each with a different size of international arbitrage costs, one would expect the adjustment of the overall real exchange rate to be smooth rather than discontinuous, as noted by Taylor et al. (2001). Teräsvirta (1994), in turn, shows that time aggregation is also likely to result in smooth regime changes rather than discrete ones as long as heterogeneous agents do not act simultaneously.

An alternative characterization of the discrete nonlinear adjustment is provided by smooth transition autoregressive models (STAR). Here, in contrast to the discrete SETAR model, regime changes occur gradually and are determined with a smooth function, which need only be continuous and non-decreasing. The assumption of heterogeneous agents and the need for symmetry in the response to positive and negative deviations from PPP leads empirical studies to the exponential STAR (ESTAR) model.

$$G(\gamma, c, z_t) = 1 - \exp\{-\gamma(z_t - c)^2\} \qquad \gamma > 0. \tag{1}$$

The transition function goes from zero to one as z_t, the transition variable, increases. The slope parameter γ indicates how rapid the transition from zero to unity is as a function of z_t. Finally, c is the location parameter, which determines where the transition occurs. Michael et al. (1997), Baum, et al. (2001) and Taylor et al. (2001), among others, apply ESTAR model and find support for the nonlinear representation.[4]

The transition function in the ESTAR model is symmetric around c and $G(\gamma, c, z_t) \rightarrow 1$ for $z_t \rightarrow \pm\infty$. A drawback of the exponential transition function is that for either $\gamma \rightarrow 0$ or $\gamma \rightarrow \infty$, the transition function collapses to a constant and the model becomes linear in both cases. Hence, the ESTAR model does not nest a SETAR model as a special case, as noted by Van Dijk et al. (2002).

An exponential transition function is a suitable transition function if we assume non-linearity in the model is due to symmetric and heterogeneous transaction costs. Similarly, if the beliefs of heterogeneous agents determine the exchange rate behavior, it is possible to model the smoothness between regimes using the ESTAR model. It is desirable, however, to allow the adjustment towards long-run equilibrium to be discrete as well as smooth if the policy rule followed by the central bank encourages market agents to enter the market at the same time. We use a second-order logistic function (LSTAR2)

$$G(\gamma, c, z_t) = (1 + \exp\{-\gamma(z_t - c_1)(z_t - c_2)\})^{-1}. \tag{2}$$

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[3] Obstfeld and Taylor (1997) use data measured relatively to the US after 1980.
[4] Michael et al. (1997) use long time series spanning 1791-1992 and also post-war time series for bilateral U.S. real exchange rates. Baum et al. (2001) examine bilateral U.S dollar CPI and WPI proxies over the post Bretton -Wood period. The data set in Taylor et al. (2001) comprises several bilateral real exchange rates against U.S. dollar during the post-Bretton-Wood period.

If $\gamma \to 0$, the model becomes linear, whereas if $\gamma \to \infty$ and $c_1 \neq c_2$, the function $G(\gamma, c, z_t)$ is equal to 1 for $z_t < c_1$ and $z_t > c_2$, and equal to 0 in between. Thus, the LSTAR2 model nests a SETAR model as a special case.

3.2 DATA

The sample consists of monthly observations from October 1982, the date of a possible regime shift in U.S. monetary policy, to June 2003.[5] The advent of the euro did not necessarily drastically change the weight put on the bilateral dollar-mark (euro) in monetary policy rules. This is because, by definition, the weight of this bilateral exchange rate has increased in monetary policy rules and also because EMU currencies were relatively stable before the advent of the euro. The source of the data was the OECD's Main Economic Indicators. The exchange rate set of variables is defined by

$ppp_t = p_t - p_t^* - s_t$, where

p_t = the German (euro), or 'home', consumer price index,

p_t^* = the US, or 'foreign', consumer price index,

s_t = the spot exchange rate, defined as \$/DM(euro),

All the variables are in logarithmic forms. Finally, the persistence of the nominal exchange rate deviations from the PPP equilibrium is apparent in Figure 1. It also appears that the persistency does not arise as the result of a single period, such as the large swings of the mid 80's.

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[5] From 1979 to 1982, the Federal Reserve (FED) targeted non-borrowed reserves. Since then, the FED has been followed a monetary policy rule which practically amounts to an interest rate targeting policy.
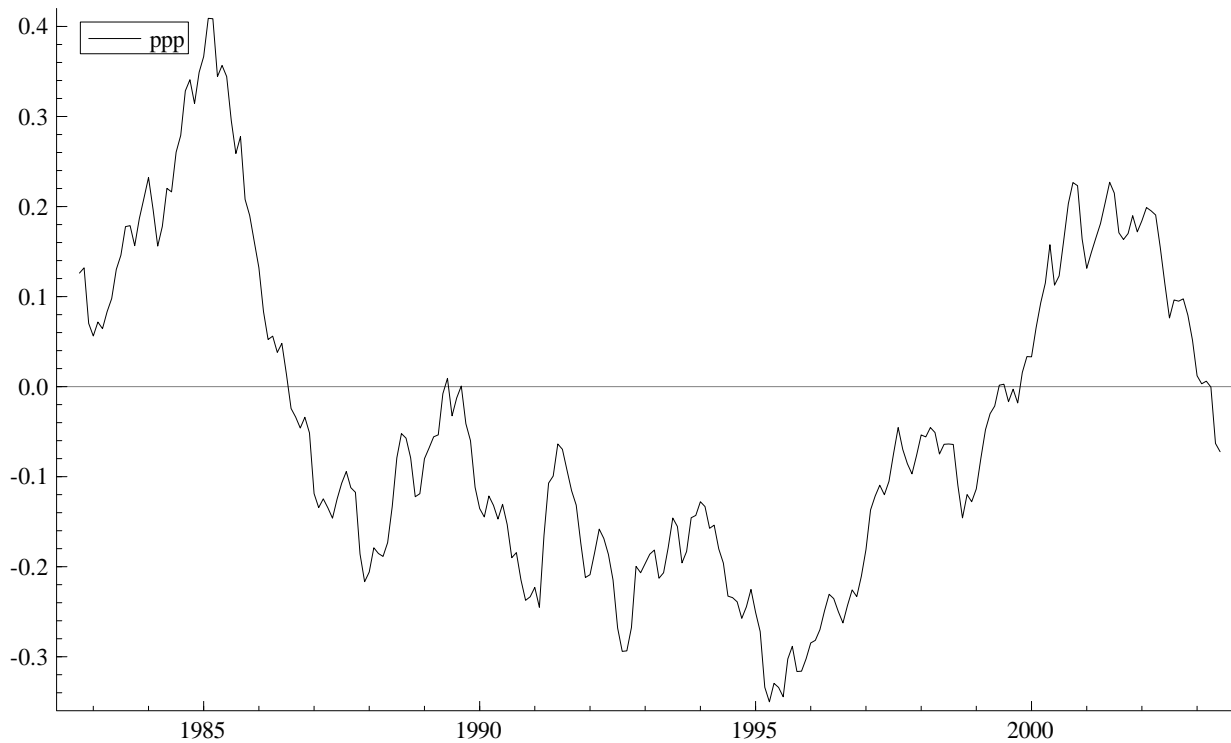

Table 1. ADF and Eklund-F tests.

82.10-2003.6	obs:	249	
PPP			
ADF(t),1	sigma		
drift	0,32	-1,82	
no drift	0,32	-1,78	*
ADF(t),1,dif			
drift	-0,73	-11,45	***
no drift	-0,75	-10,03	***
Eklund(F)			
drift	0,36	1,35	
no drift	0,36	1,14	
* = rejection of the null hypothesis at the 10 % level of significance			
**= 5%			
***=1%			

We use the Eklund-F test (EF) for the joint testing of linearity and unit root hypothesis against the second-order logistic STAR (Eklund, 2003). The artificial regression is as follows:

$$y_t = \delta_1 \Delta y_{t-1} + \delta_2 (\Delta y_{t-1})^2 + \delta_3 (\Delta y_{t-1})^3 +$$
$$\phi_1\ y_{t-1}\Delta y_{t-1} + \phi_2\ y_{t-1}(\Delta y_{t-1})^2 + \alpha + \rho y_{t-1} + \varepsilon_t \tag{3}$$

A joint test of linearity and the unit root hypothesis amounts to testing the hypothesis $H_{01}: \delta_2 = \delta_3 = \phi_1 = \phi_2 = \alpha = 0, \rho = 1$ in (3). The null hypothesis is that the true data generating process is a random walk. Excluding $\alpha = 0$ from H_{01} results in another null hypothesis H_{02} that allows for a unit root process with drift component. Since under the null hypothesis the real exchange rate follows a unit root process, the null hypothesis complicates the testing procedure analogously to the way in which the distribution of a Dickey-Fuller statistic cannot be assumed to be Students's t. A Monte Carlo study of the critical values, size and power properties of the EF is provided by Eklund (2003).[7] The rejection of the null hypothesis provides evidence of stationary but nonlinear time series.[8] The results do not, however, give any support to the conclusion that the real exchange rate follows a stationary and nonlinear process.

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[6] The results did not change if a linear trend was included in the Dickey-Fuller regression.

[7] According to Eklund (2003) the size of the test is distorted when the value of δ_1 is close to –1 or 1. The size of the ADF test is distorted only when δ_1 is close to 1.

[8] We use critical values for 250 observations.

The ADF tests lack power against the stationary PPP alternative over the post-Bretton-Wood time period. Taylor et al. (2001) pointed out that the unit root behavior does not necessarily imply that no long-run equilibrium exists. The failure to reject a unit root may indicate, conversely, that most of the time real exchange rates have been in the neighborhood of the long-run equilibrium level. This is because real exchange rates behave more like unit root processes the closer they are to long-run equilibrium. We call this feature of real exchange rate time series local unit root behavior.

Although the power simulations in Eklund (2003) show some gain in power compared to the standard ADF test, EF also has low power in discriminating a random walk from a theoretically meaningful stationary and nonlinear PPP alternative at the sample size available for the tests. Thus, we build a univariate model for the U.S. dollar- German mark (euro) real exchange rate to find out whether local unit root behavior dominates during the time period 1982:10-2003:6.

3.4. UNIVARIATE MODEL FOR THE REAL EXCHANGE RATE

We begin our analysis by assuming that the real exchange rate series is stationary. We also assume linearity, but we prepare to consider the possibility that the real exchange rate cannot be adequately characterized by a linear autoregressive model. Our alternative model to the linear model is the smooth autoregressive model (STAR)

$$q_t = \phi' x_t + \theta' x_t G(\gamma, c, z_t) + \mu_t, \qquad (4)$$

$x_t = (1, q_{t-1}, ..., q_{t-p})'$, $\phi = (\phi_0, \phi_1, ..., \phi_p)'$, $\theta = (\theta_0, \theta_1, ..., \theta_p)'$, and $\mu_t \sim nid(0, \sigma^2)$.

where G is a transition function.

The first step is to specify a linear AR(p) model for the real exchange rate series to serve as our null hypothesis. The order of autoregression, p=2, is chosen through inspection of the partial autocorrelation function. Table 2 presents the results of testing for nonlinearity using the third order

artificial regression suggested by Luukkonen et al. (1998). The delay length, d, is varied in order to provide the strongest probability of non-linearity.[9]

Table 2. Linearity against non-linearity.

	p-Values of Tests of Linearity against STR											
	LAG											
PPP	1	2	3	4	5	6	7	8	9	10	11	12
F	0,32	0,38	0,37	0,25	0,22	0,12	0,16	0,10	0,03	0,04	0,06	0,07
F(4)									0,13	0,10		
F(3)									0,01	0,02		
F(2)									0,97	0,83		

The null of linearity is strongly rejected when p= 2 and d = 9 or 10. The choice between transition functions is based on the test sequence suggested by Teräsvirta (1994).[10] This supports either an LSTAR2 or an ESTAR function. We will use an LSTAR2 model. The final choice between them was made based in the discussion in Chapter 2.

We subsequently found that the simplifying restriction $\phi_2 = 1 - \phi_1$ could not be rejected at standard significance levels for any of the estimates. Furthermore, the nonlinear part of autoregressive parameters always provides negative values. This gives the model a local unit root when G = 0.

We use PPP_10 (lag 10) as a transition variable. The final choice between PPP_9 and PPP_10 was based on econometric considerations, i.e. we are not able to calculate standard deviations of Gamma and c_1 variables if PPP_9 is used as a transition variable. Since the PPP_10 provides more accurate parameter estimates in the following analysis, it is assumed that the data has been generated by such a LSTR2 model.

Table 3 shows the estimated results. Bold values correspond to a significant value at the 5% level. The figures in the last column are the percentages of observations below c1 and above c2 respectively.

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[9] We use ordinary F-test, since, as found by Granger and Teräsvirta (1993), an F-approximation works much better with small sample size than LM test with the asymptotic χ^2 distribution.

[10] The p-values for the whole sequence of test are given only if the general linearity test (F) lies below 0.05.

Table 3. Nonlinear model for real exchange rate.

Linear	drift	PPP_1	Non-lin.	Gamma	Low	High	drift	PPP_1	%
	0,00	1,29		211,39	-0,20	0,19	-0,01	-0,07	31

$R^2 = 0,98$, $\hat{\sigma} = 0,02$, $\hat{\sigma}/\hat{\sigma}_l = 0,96$, sk = -0,23, k =3,37

$\hat{\sigma}$ denotes the residual standard deviation of the nonlinear model. The residual standard deviation of the nonlinear model is about 96 per cent of that of the corresponding linear model. This high correlation occurs because most of the time the real exchange rate fluctuations fall within the discrete borders. Note also that we included a drift term to capture difficulties in identifying the long run equilibrium level of the real exchange rate.

Residual diagnostic tests are reported in Table 4. Tests of no error autocorrelation are Lagrange multiplier test statistics for the first order up to sixth order serial correlation in the residual, as discussed in Eithreim and Teräsvirta (1996). A traditional ARCH LM test is used for autoregressive conditional heteroskedasticity. The results of the LM test of no error autocorrelation do not indicate autocorrelation nor is there any evidence of ARCH. Thus, the LSTR2 model appears to provide an acceptable representation for the adjustment process towards PPP.

Table 4. Residual diagnostic tests.

	P-values of the LM test of no Error Autocorrelation against AR(p) and MA(p) Error Process, and the LM- test of no Autoregressive Condotional Heteroskedasticity against ARCH(p). Maximum Lag p					
Test	1	2	3	4	5	6
No error autocorrelation	0,07	0,15	0,23	0,31	0,44	0,37
No ARCH	0,42	0,71	0,72	0,58	0,60	0,49

A natural question is whether the rejection of the linear model is perhaps driven by a single episode in the data. Probably this is not the case since there are quite many observations outside the band (31%). Gamma is adjusted by $\sigma^2(ppp_t)$ which is the sample variance of ppp_t. Since $\gamma \to \infty$ then $G(\gamma,c,z_t) \to 0$ for $C1 \le z_t \le C2$; and for other values $G(\gamma,c,z_t) \to 1$.

In our view, our model captures two essential features of the data: nonlinear deviations from PPP, and a discrete band around PPP. The main difference in the fits of the two models (linear/nonlinear) is due to different characterization of the large fluctuations. Figure 2 shows that the transition function obtains a value of one especially during the dollar appreciation period in the eighties and again during the dollar depreciation period in the middle of the nineties.
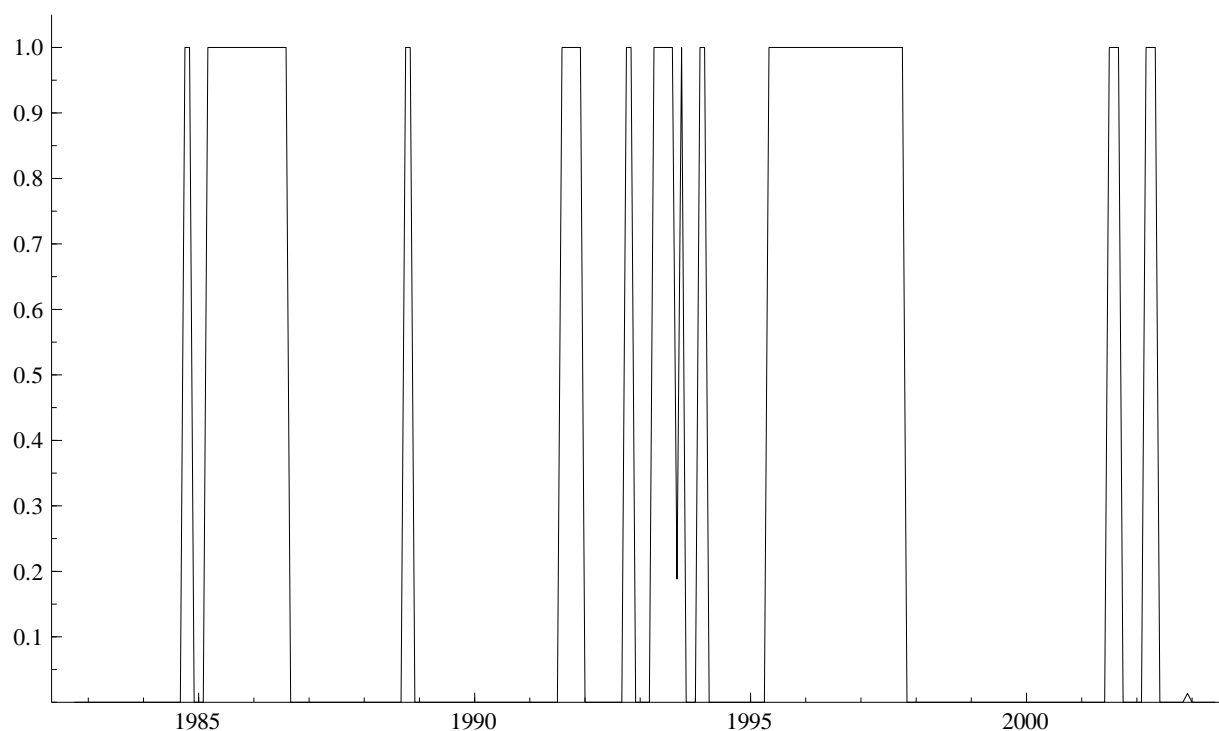
Figure 2. Values of the transition function.

The real exchange rate time series seems to follow local unit root behavior when the transition variable is not large. The failure to reject a unit root seems to indicate that the real exchange rate has most of the time been relatively close to equilibrium, rather than that the real exchange rate is not a globally stationary process. Thus, our results indicate the presence of nonlinear adjustment towards parity level.

Although we do not explicitly prove our stationarity assumption, the empirical results shown in Table 3 reveal under the assumption of stationarity that the estimated transition parameters are highly significantly different from zero, which itself indicates a correct specification. Combining this evidence with the strong theoretical and empirical evidence for nonlinear real exchange rate, makes a stationarity assumption an interesting alternative to a non-stationarity assumption based on linear unit root tests.

Our model is very close to a switching regression model.[11] This contrasts with the results that one would expect considering the ESTAR transition function, as several authors have assumed. Notably, this characteristic may imply that the central bank(s) has a role in determining exchange rates as discussed in Chapter 2.

4. CONCLUSIONS

The band around PPP is wide. This implies that the weight put on the real exchange rate is quite low. The bilateral U.S. dollar -German mark (euro) real exchange rate is also probably more important variable in the Bundesbank (the European Central Bank) monetary rule than in the Federal Reserve monetary rule. Clarida et al. (1998) estimate linear monetary policy reaction functions for Germany and Japan. They find that a one percent depreciation of the mark relative to the dollar led the Bundesbank to increase interest rates by five basis points. It has been very difficult to find such an effect for the United States.

Our model is also very close to a switching regression model. This may indicate that near the lower and upper boundaries the expectations on future monetary reactions are increased, i.e. traders have homogeneous expectations near the boundaries. This is in all probability due to official announcements by the central bank(s).

The values of lower and upper boundaries are realistic. We find, however, that the order of delay parameter is nine or ten. Although large d values are also typical findings in ESTAR models, they are problematic if we do not have any clear economic intuition to support large d values.[12] In our model this may imply a lack of policy co-ordination, which creates temporary speculative bubbles. This points to a need for further research.

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[11] See a similar finding in Bec et al. (2002). They use a three-regime LSTAR model with the symmetry restriction.
[12] Taylor et al. (2001) is an exception (d =1). For an example of large d value, see Baum et al. (2001).

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