Nonlinearities in International Macroeconomics: An empirical analysis of advanced economies and emerging markets

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I dedicate this thesis to my parents, for their love and support.

I remain responsible for any remaining errors and for the views expressed in this thesis.

Abstract

The empirical literature in international macroeconomics and finance reveals a growing interest in the use of nonlinear models. Their attractiveness is clear, as these models allow capturing discontinuity in the data-generating process and, by estimating endogenously transition probabilities and variables, avoid the *a priori* identification of regimes and their timing, thereby enhancing the extent of flexibility in the analysis. This thesis makes use of nonlinear techniques in order to model two different economic issues, which have been at the centre of the economic debate in the last years.

The first analysis refers to the issue of debt sustainability and, in particular, tries to test empirically some of the leading interpretations that have been advanced to account for the financial turmoil that characterised the run-up to the Brazilian presidential elections in 2002. We test for financial contagion from the Argentine crisis and the impact of factors including IMF intervention and political uncertainty in raising the probability of crisis. The empirical investigation employs a Markov-switching model with endogenous transition probabilities.

The second part of the thesis is devoted to the analysis of current account imbalances in G7 countries. We find evidence of threshold behaviour in current account adjustment, such that the dynamics of adjustment towards equilibrium depend upon whether the current-account/ net-output ratio breaches estimated country specific current account surplus or deficit thresholds. Both the speeds of adjustment and the size of the thresholds are found to differ significantly across countries. We complement the univariate analysis by disentangling the domestic components of the current account according to the national income identity with a view to shed light on the role of savings (both public and private) and investment. Evidence of shifts in means and variances of exchange rate changes - that coincide with the current account adjustment regimes identified by the model - suggest scope for further research on the role of the real exchange rate in determining the nonlinear behaviour of the current account. We extend the threshold methodology to a bivariate context and find evidence of a strict link between current account adjustments and deviations of a country real exchange rate from its long run equilibrium, such that beyond a certain appreciation/depreciation of the real exchange rate, a country CA imbalance would start reverting towards its mean value. Finally, we run a nonlinear test of the present value model of the current account, encouraged by the evidence of nonlinear adjustment in the current account. However, in line with the literature, we reject the test for all countries also in this new nonlinear framework.

Chapter 1
Introduction

The use of nonlinear models in time series analysis has become increasingly popular over the last decade. Their attractiveness in the analysis of issues in international macroeconomics and finance is clear, as these models allow to capture discontinuity in the data-generating process, in the form of regime changes due to a structural alteration in the economic system as well as extraordinary short-period events such as a financial crisis. Moreover, by estimating endogenously transition probabilities and variables, nonlinear models avoid the *a priori* identification of regimes and their timing, thereby enhancing the extent of flexibility in the analysis. Given these attractive features, it is not surprising that the empirical literature in international macroeconomics reveals a growing interest in the use of this category of models.

This thesis makes use of nonlinear techniques in order to model two different economic issues, which have been at the centre of the economic debate in the last years.

The first analysis refers to the issue of debt sustainability in emerging markets and, in particular, tries to test empirically some of the leading interpretations that have been advanced to account for the Brazilian financial turmoil in 2002. In the run-up to the presidential elections, a strong currency depreciation and a sudden and sharp rise in sovereign

bond spreads put Brazil under strain, by severely impacting upon its indexed debt, despite the important reforms of the previous years and the relative sound fundamentals.

The nature of the topic requires a methodology able to account for the existence of unobservable factors (such as changes in investors' beliefs, herding behaviour, financial panic and political uncertainty) that are likely to affect the process under investigation. Markov-switching (MS) models with endogenous transition probabilities prove a valuable tool for the purposes of modeling regime shifts. In effect, the possibility of capturing self-fulfilling changes in market behaviour, by means of an endogenous regime-shift selection, allows researchers to test multiple equilibria models empirically.

The fact that the crisis occurred during the electoral campaign, coupled with the concerns expressed by some financial commentators relating to the expected victory of the socialist candidate, Lula da Silva, a former trade unionist known for his past declarations on debt repudiation, led several authors to interpret the Brazilian experience as a financial distress episode triggered by uncertainty on the political outcome. In the analysis we therefore assess the existence of a political "mistrust effect" for the left-wing candidate on the country default risk by means of an opinion polls variable. In light of the literature on catalytic finance, we also address the role of IMF intervention on investor sentiment, by creating an IMF dummy variable including both positive and negative news from the IMF or domestic politicians. We test for both financial and political contagion between Argentina and Brazil by extending the analysis to Argentine data. Finally, we account for a global factor effect by introducing into the empirical model the high-yield spread in developed markets.

The second part of the thesis is devoted to the analysis of sustainability and adjustment of current account imbalances in G7 countries, a subject that has been receiving considerable attention among policymakers, financial market practitioners and academics. At more than 6 percent of GDP, the US current account deficit attracts the most focus, but there are also material current account imbalances in other deficit countries such as the UK and in surplus countries such as Japan and Germany. The question at the heart of the present analysis is whether or not the stationary stochastic process which describes the current account to net output ratio in the G7 countries features linear or nonlinear adjustment to the unconditional mean. We identify threshold autoregressive models as tractable and testable nonlinear time series models that conveniently exhibit all of the features of the current account adjustment process that have been the focus of recent discussions in the literature, and that nest as a special case the linear stationary stochastic process model for the current account that is often assumed in empirical work.

In Chapter 3, we estimate for each G7 country an univariate threshold autoregressive model of the current account to net output ratio, allowing for country-specific thresholds of current account surplus and deficit adjustment in each country, and also allowing for country specific means for the ratio of the current account to net output. We then investigate what happens to the probability distributions of nominal exchange rate changes, stock price index changes, and long term interest rate differentials during the various current account adjustment regimes. The motivation is to determine whether or not crossing the current account adjustment threshold is itself associated with shifts in the probability distributions for exchange rates, stock prices, and interest differentials. Finally, we draw on our empirical

results to take stock of the present US current account deficit, characterised by relatively wide thresholds within which current account adjustment is absent and by relatively slow speeds of adjustment once these thresholds, especially the deficit threshold, are crossed. We explore possible unusual circumstances, which might have delayed the US current account adjustment and that were not in evidence during the sample over which the models were estimated, 1979-2003.

Chapter 4 complements the univariate results on current account adjustment by providing a disaggregated nonlinear analysis of its domestic components according to the national income identity. After investigating stilysed facts on the current degree of international capital mobility and the robustness of the assumption of Ricardian equivalence, we highlight the different role of savings and investment on current account adjustments in G7 countries.

In Chapter 5, we attempt at providing a comprehensive explanation of the joint dynamics of real exchange rate and current account, in particular of the role of the real exchange rate in determining the nonlinear adjustment of the current account. Our analysis builds on the vast literature on the existence of nonlinearities in the dynamics of the real exchange rate as well as on the preliminary evidence provided in chapter 3 of statistically significant shifts in the mean and variance of the probability distribution of exchange rates that occurred in conjunction with current account adjustment regimes. By extending the threshold autoregressive methodology to a bivariate context, we find evidence of a strict link between current account adjustments and deviations of a country real exchange rate from its long run equilibrium, such that beyond a certain appreciation/depreciation level

of the real exchange rate, a country CA imbalance would start reverting towards its mean value.

Chapter 6 reports research on a nonlinear test of the present value model of the current account. We estimate a multivariate system for current account, net output, world real interest rate and real exchange rate, allowing for a threshold adjustment in the current account and exchange rate variables, and compute forecasts of these variables over a finite sample, as suggested by the speed of adjustment of their temporary components. The evidence on the threshold behaviour of the current account and on the significant effect of the real exchange rate on its adjustment, as presented in the previous chapters, seemed to support this type of methodology. Nevertheless, in our analysis of the intertemporal model of the current account, the predictive power of our forecasts remain highly unsatisfactory.

Chapter 2
The Brazilian Currency Turmoil of 2002: A Nonlinear Analysis

2.1 Introduction

During the last decade or so, a number of Latin American countries experienced severe financial and currency crises, from Mexico in 1994 to Argentina in 2001. Debt sustainability problems have been a crucial issue in all of these episodes of distress. Even when a country's foreign debt exposure was relatively low, the structure of its sovereign bonds made it vulnerable to speculative attack or any other financial market turbulence. In the spring of 2002, a strong currency depreciation and a sudden and sharp rise in sovereign bond spreads put Brazil under strain, by severely impacting upon its indexed debt, despite the important reforms of the previous years and the relative sound fundamentals.

The financial turmoil generated by the change in the economic environment appeared to be the result of a shift in market equilibria, from stability to crisis. The fact that the crisis occurred in the run-up to the Brazilian presidential elections in October 2002, coupled with the concerns expressed by some financial commentators relating to the expected victory of the socialist candidate, Lula da Silva, a former trade unionist known for his past declarations on debt repudiation, has led authors such as Williamson (2002) to interpret the Brazilian experience as a financial distress episode triggered by uncertainty on the political outcome. In contrast, very little attention has been devoted to the contemporaneous crisis

in Argentina as a possible explanation of the Brazilian turmoil. The Argentine debt crisis has been mainly regarded as an entirely predictable and independent event, unable to influence the other economies of the region apart from marginal financial spillovers[1]. Some authors – e.g. Miller et al. (2003) – have, however, suggested a role for the Argentine crisis and, more generally, of its underlying political events in coordinating public expectations concerning the behavior of a possible left-wing government in Brazil.

This chapter aims to investigate the events of 2002 in Brazil, by testing empirically some of the interpretations recently proposed by the theoretical literature on the subject. In light of the literature on catalytic finance and recent work by Corsetti et al. (2003) and Morris and Shin (2003), we also address the role of IMF intervention on investor sentiment. Finally, we consider the effect of global factors in the Brazilian turmoil. In particular, we look for evidence of a contemporaneous increase in global risk aversion as a potential threat to the vulnerable Brazilian economy during this period.

Building on previous work on contagion by Jeanne and Masson (1998) and Fratzscher (2002) as well as the econometric literature on business cycles (Filardo and Gordon, 1998), we apply a time-varying transition probability Markov-switching model to the analysis. We find the use of such methodology particularly useful in detecting shifts in regime and find that its performance is superior to simple linear models. In contrast to previous work on this issue, we employ high-frequency financial data, as they seem better able to capture shifts in investors' behaviour. Specifically, we use sovereign bond spreads as a proxy for country default risk. We assess the existence of a political "mistrust effect" for the left-wing

[1] See Krueger [2002].

candidate by means of an opinion polls variable and create an IMF dummy variable in order to capture any catalytic effect arising from both positive and negative news from the IMF or domestic politicians. We test for both financial and political contagion between Argentina and Brazil by extending the analysis to Argentine data. Finally, we account for a global factor effect by introducing into the empirical model the high-yield spread in developed markets.

The empirical results strongly support the hypothesis of a shift in regimes in Brazil during 2002, and indicate the political instability of the pre-electoral period and the global increase in market risk aversion as main causes of the jump in equilibria. Negative news concerning IMF intervention also has a significant impact on the probability of shifting between regimes. Argentine country risk, however, affects only the *extent* of the Brazilian turmoil, without triggering it.

The remainder of the chapter is organised as follows. Section II reviews the main events in Brazil during the summer of 2002 and the problem of public debt sustainability. Section III addresses the main interpretations of the 2002 events with reference to the current economic debate and the recent literature on financial crises. Section IV presents the Markov-switching methodology adopted in the analysis. Section V discusses the main empirical results of the model. Section VI concludes.

2.2 The Depreciation of the Brazilian Real of 2002 and the Problem of Debt Sustainability

In the summer of 2002, the Brazilian economy was subject to major currency turmoil with a depreciation of its currency, the *real*, of over 30% between April and October (see Figure. 2.1). The exchange rate disorder was associated with a net drop in capital inflows, a sharp rise in the interest rate spreads of Brazilian debt over US Treasury securities, and a fall in domestic debt rollovers. This turbulence has been interpreted as the result of financial market fears that Brazil could default on its public debt, following the example of other emerging markets and especially that of its closest neighbour, Argentina.

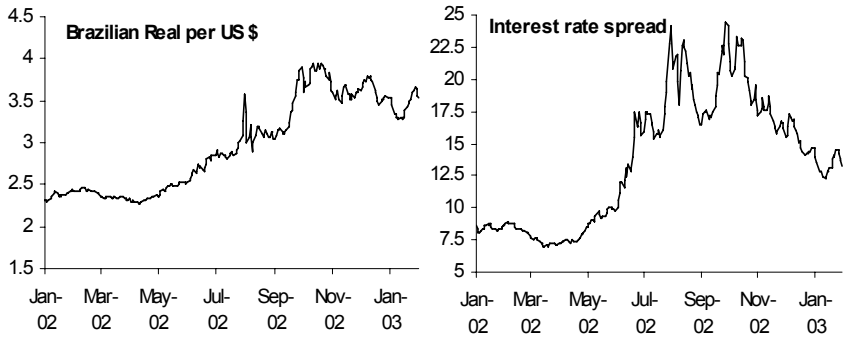

Figure 2.1: Nominal Exchange Rate and Bond Spreads

Sources: DataStream, JP Morgan (EMBI Global)

In the years preceding this financial distress episode, however, the Brazilian authorities had achieved a substantial improvement in the country's institutional framework: the 1998 Fiscal Stabilisation Plan[2] and the 2000 Fiscal Responsibility Law—as well as the

[2] This law governs the spending patterns of the country's federal states, by establishing limits on public

greater transparency that came with the publication of improved government statistics and the recognition of fiscal hidden liabilities—being among the most relevant reforms[3]. Since 1995, the fiscal balances had registered significant primary surpluses averaging about 3.5% of GDP and high levels of fiscal revenue had been collected. The macroeconomic environment also benefited from an effective inflation-targeting policy, which delivered relatively stable inflation in spite of the large currency depreciation. With respect to the trade position, the competitive exchange rate allowed an improvement of the current account balance (Figure 2.2). These outcomes are even more remarkable in light of the deteriorating macroeconomic situation in neighbouring Argentina, Brazil's single-largest trading partner[4]: it is, in fact, estimated that Brazilian exports to new markets were able to offset more than 80% of the drop in exports to Argentina.

This economic framework seems at first sight to contain all the necessary ingredients for a successful trend in Brazilian fundamentals, specifically with regard to debt dynamics. On the basis of these results and of a consistent medium and long-term scenario, an IMF fiscal sustainability analysis, released in January 2001, estimated a gradual net public debt decline over time, assuming a continuation in the fiscal efforts of the federal government.

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indebtedness or expenditure in personnel, and defining annual fiscal targets for three successive years.

[3] See Goldfajn [2002].

[4] The weight of the Argentine peso in the real effective exchange rate calculated by Banco do Brasil is equal to 15.8% (second only to the US dollar with 32.4%) and reflects trade during the period 1998-2001.

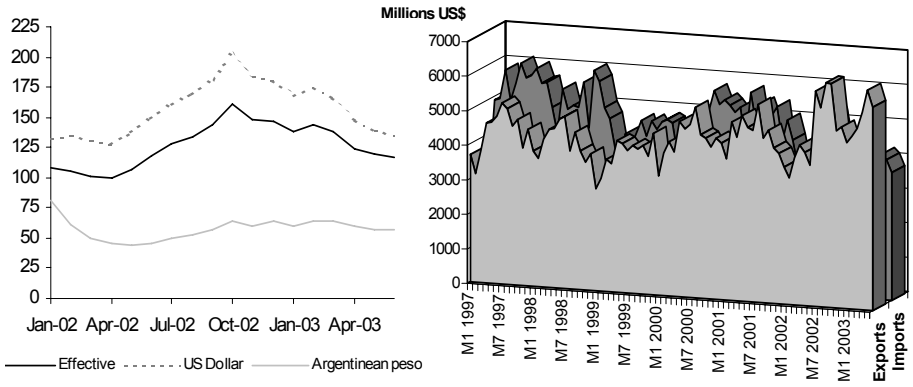


Figure 2.2: Real exchange rates and trade dynamics

Sources: Banco Central do Brasil, IFS

Nevertheless, these positive signals were unable to quell investors' concerns over the sustainability of Brazilian debt. In fact, the vulnerability of the Brazilian economy to shifts in investor sentiment and other external shocks depended critically on the composition of its public sector debt. Increased volatility in emerging markets in previous years, coupled with Brazil's past history of monetary and fiscal mismanagement, led to a strong demand for indexed debt from foreign as well as domestic investors. As a result, beginning with the East Asian crisis in 1997, a large share of prefixed Brazilian public debt had been converted into indexed debt. As depicted in Figure 2.3, the share of public debt with a floating interest rate[5] (37% of the total) or linked to the exchange rate (42%) represented a major source of vulnerability for the country. As a result, panic behaviour in the summer of 2002 was more than sufficient to offset the efforts of the Brazilian authorities: any hostile market

[5] The Selic is the overnight interest rate that is set by the Brazilian Central Bank.

sentiment was self-validated by the rise in bond spreads and depreciation of the currency, which pushed the debt-to-GDP ratio up to around 56% by end-2002[6].

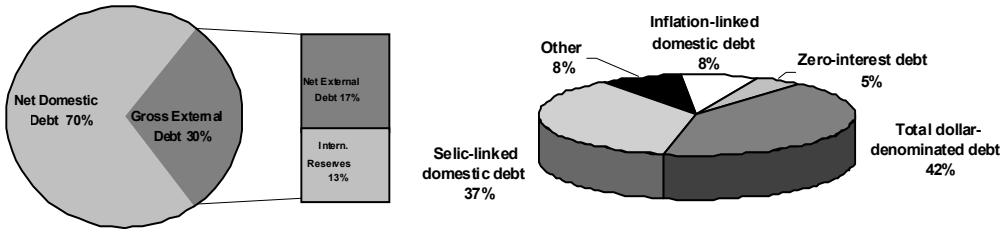


Figure 2.3: Composition of Brazilian Public-Sector Total Debt (Aug-2002)

Sources: Williamson [2002], Banco Central do Brasil.

2.3 Financial Contagion and Political Uncertainty

2.3.1 A Multiple Equilibria Story?

The 2002 currency panic in many ways resembles a self-fulfilling crisis episode, as opposed to a "wake-up call" effect, where an external event reveals to investors the true state of the fundamentals. In effect, the fiscal situation of Brazil in 2001 was far from unsustainable; yet the level of public debt and, in particular, its flexible structure were not adequate to keep the economy in the no-crisis state and left it vulnerable to shifts in investors' expectations.

[6] Williamson [2002] regards the 56% figure declared by the Brazilian CB as too low and takes this figure up to 66%. Goldstein [2003] reports similar estimates for major rating agency and investment banks. Favero and Giavazzi [2002] remark on the existence of hidden liabilities that are still not officially recognised and that may account for as much as 10% of GDP.

As described by second-generation crisis models,[7] when an economy is exposed to multiple equilibria, sunspots—i.e. factors able to coordinate and redirect investors' expectations—can induce jumps between them. Such a mechanism is enabled only within a certain range of the fundamentals necessary to place the country in the crisis zone: low reserves, a negative trade balance and/or a large variable debt, as in the Brazilian case, are among the most likely causes of vulnerability. According to this theoretical framework, it follows that, given its fiscal situation, Brazil had two alternative stable equilibria: a "good" equilibrium, characterised by constant capital inflows, relatively low interest rates, a stabilisation of the exchange rate and, as a result, a manageable fiscal burden; and a "bad" equilibrium, characterised by diminished foreign investment, high interest rates, an over-depreciated currency and a far from sustainable debt-service figure.

Nevertheless, recent work on currency crises by Morris and Shin (1998) has focused attention on the role of uncertainty and lack of common knowledge. In these models, the introduction of noisy private information about fundamentals eliminates any common device able to coordinate investors' actions and leads eventually, by iterated elimination in a global-game framework, to selection of a unique equilibrium. Morris and Shin (2001) (and Prati and Sbracia, 2003) extend this analysis in order to distinguish between different sources of uncertainty (fundamental and strategic) as well as a different taxonomy of information (private or publicly available). The result is a condition for uniqueness of equilibrium which relies strongly on the precision of private signals relative to the underlying uncertainty on the fundamentals. From this perspective, the Brazilian events of 2002 may

[7] See the seminal paper of second-generation crisis models by Obstfeld (1996) or, for a text-book treatment, Sarno and Taylor (2003).

have been caused by a shift in fundamentals, where each agent takes a global perspective and deduces when an adequate number of other players is ready to attack the currency. This scenario needs by assumption to focus on the noisy character of the surrounding public signals. In particular, with reference to the presidential campaign in Brazil during 2002, uncertainty over economic fundamentals could include uncertainty over "political fundamentals", assessed by agents according to their correlated private signals together with the noisy public signals in the form of public opinion polls.

Notice, however, that if the strength and direction of common signals overcame the precision of private information, we would be brought back to a multiple-equilibria scenario,[8] where sunspots are replaced by increases in the noise of individuals' signals.[9] Following this latter interpretation, the precision of private information may have been further compromised by the vulnerable character of indexed Brazilian debt, since under indexation a rise in the sovereign spread will actually increase the size of the debt, which may in turn adversely affect market sentiment in a vicious spiral.

2.4 Modelling Crisis in a Nonlinear Framework

2.4.1 Nonlinear Analysis of International Financial Markets: A Brief Review

The use of linear models in the analysis of financial time series presents a number of limitations, as their formulation does not allow for the existence of unobservable factors (such as

[8] See Hellwig (2002).

[9] Heinemann and Illing (2002).

changes in investors' beliefs, herding behaviour, financial panic and political uncertainty) that are likely to affect the process under investigation. For this reason, nonlinear empirical analyses of financial and international financial variables have become increasingly popular over the last decade or so. In particular, Markov-switching (MS) models represent a valuable tool for the purposes of the present analysis.[10] Their attractiveness in the context of the present analysis are clear: As MS models are designed explicitly to capture discontinuity in the data-generating process, they allow us to model regime changes due to a structural alteration in the economic system as well as extraordinary short-period events such as a financial turmoil episode. Moreover, by allowing the estimation method to determine the probability of being in any particular regime, we can avoid the *a priori* identification of regimes and their timing, thereby enhancing the extent of flexibility in the analysis.

Given these attractive features, it is not surprising that the empirical literature on financial crises and contagion reveals a growing interest in the use of this category of models. The work of Jeanne and Masson (1998), who propose the utilization of a MS framework in order to depict empirically the existence of multiple equilibria, represents one of the first attempts in this vein. Specifically, these authors test a model of self-fulfilling expectations in the speculation episode against the French franc in the period 1992-1993. Their results show that the performance of the model improves significantly once sunspots are introduced to influence devaluation expectations, by means of a MS approach. In Masson (1999), however, some limitations in the use of MSVAR to the analysis of contagion are discussed: in particular, the application of MS analysis to interest rates for Argentina

[10] For a wide coverage of Markov switching models refer to Kim and Nelson [1999].

and Brazil seems not to produce the expected results. Among the main problems identified by Masson is the availability of an adequate data sample: the use of high frequency data and the analysis of large asset price movements rather than crises per se are suggested as possible alternatives. Fratzscher (2002) employs a MSVAR methodology to allow for a systematic comparison of the possible explanations of financial crises: fundamentals, sunspots and contagion. Fratzscher's empirical study is based on monthly data from 24 emerging markets over the period 1986-1998. Only the inclusion of contagion variables—measuring trade, financial and stock market interdependence across countries–allows the linear model to perform as well as the nonlinear one, suggesting the effectiveness of MS modelling in the detection of latent, crisis-generating factors. The MS methodology is applied to high-frequency data by Sola et al. (2002) in order to account for changes in expectations and investors' beliefs in the test of volatility spillovers across stock markets. Finally, Tillmann (2004) adopts a time-varying regime-switching regression in order to model the probability of a currency crisis for the French franc and the Italian lira under the ERM. In particular, the model tests for the theoretical predictions of models *à la* Morris and Shin (1998), by introducing a measure of information disparities.

2.4.2 A Markov-Switching Model of Crisis with Endogenous Transition Probabilities

In the present analysis, we use the Brazilian sovereign bond spread as a proxy for the perceived risk of default, and estimate an equation explaining the behaviour of this equation

that has a Markov-switching mean, in order to capture the presence of an immediate jump of the series vector to its new level[11]. The model we consider is therefore of the form:

$$BRSPR_t = \alpha_0 + \alpha_1 S_t + \beta' x_t + \varepsilon_t, \quad \varepsilon_t \tilde{} N(0, \sigma^2), \tag{2.1}$$

where $BRSPR_t$ denotes the Brazilian spread, α_0 and α_1 are scalar parameters, x_t is a vector of variables that influence the level of the spread, β is an associated vector of coefficients and ε_t is a white-noise disturbance term. In this setup, a shift into crisis is represented by a shift from the 'normal state', where $S_t = 0$ and the intercept term is just α_0, to the crisis state, where $S_t = 1$ and the intercept term becomes $\alpha_0 + \alpha_1$ so that (assuming $\alpha_1 > 0$) a higher level of the sovereign spread is indicated, for any given values of x_t.

In the standard MS model, the probabilities of switching between regimes are assumed to be constant and exogenous. In the present analysis, however, we want to extend this approach in order to examine whether the probability of being in a crisis state is dependent upon one or more of a range of variables, as discussed above. Following Filardo and Gordon (1998), we can model these probabilities using a latent variable probit model, such that

$$\Pr\{S_t = 1\} = \Pr\{S_t^* \geq 0\}, \tag{2.2}$$

where S_t^* is a latent variable defined by the following equation

[11] The same indicators estimated in a simple linear framework are not significant, with the only exception of the Argentine spreads. The variables of the linear model, whose results we omit for brevity, are estimated in differences, as integrated of order one and not cointegrated. However, the same variables are found stationary at 10% critical level once structural breaks are accounted for, following the testing methodology by Perron (1990). This result justifies the use of levels in the Markov switching specification.

$$S_t^* = \gamma_0 + \gamma_1 S_{t-1} + \delta' z_t + u_t, \tag{2.3}$$

in which z_t is a vector of variables that influence the transition probability with corresponding factor loadings determined by the δ vector of parameters, γ_0 and γ_1 are scalar parameters and u_t is a standard normally distributed white-noise disturbance.[12] The transition probabilities can then be derived by evaluating the conditional cumulative distribution function for u_t. Specifically, if the probability of the economy remaining in the crisis state at time t once it is in crisis ($S_{t-1} = 1$), given the values of z_t, is p_t^{cc}, then

$$\begin{aligned} p_t^{cc} &= \Pr\{S_t = 1 | S_{t-1} = 1, z_t\} = \Pr\{u_t \geq -\gamma_0 - \gamma_1 - \delta' z_t\} \\ &= 1 - \Phi[-\gamma_0 - \gamma_1 - \delta' z_t] \end{aligned} \tag{2.4}$$

where $\Phi[.]$ denotes the standard normal cumulative density function. Similarly, the probability of remaining in the tranquil state at time t, given z_t (and, of course, $S_{t-1} = 0$) may be written:

$$\begin{aligned} p_t^{\tau\tau} &= \Pr\{S_t = 0 | S_{t-1} = 0, z_t\} = \Pr\{u_t < -\gamma_0 - \delta' z_t\} \\ &= \Phi[-\gamma_0 - \delta' z_t] \end{aligned} \tag{2.5}$$

Note that the probability of switching from state i to state j ($i, j = c, \tau$, where c denotes the crisis state and τ denotes the tranquil state) is straightforwardly given by $p_t^{ij} = 1 - p_t^{ii}$ ($i, j = c, \tau$).

[12] Normalising the variance of u_t is an identifying assumption that can be imposed without loss of generality.

Estimation of the time-invariant parameter vector $(\alpha_0, \alpha_1, \beta', \gamma_0, \gamma_1, \delta', \sigma^2)'$, together with estimated time series for the unobservable S_t, S_t^*, p_t^{11} and p_t^{22}, can be carried out in a Bayesian context using an application of the Gibbs sampler,[13] as suggested by Filardo and Gordon (1998). We followed the Gibbs-sampling methodology, although we employ diffuse priors for all of the parameters so that the resulting estimator is in fact equivalent to a standard maximum likelihood estimator.[14]

Of course, application of the model requires selecting a list of candidate variables to include in x_t and z_t. Given the previous discussion, we considered the developed market high-yield as a proxy for global factors, the Argentine sovereign spread as a potential indicator of the contagion effect, opinion polls relating to the popularity of the left-wing candidate in the 2002 Brazilian presidential elections, Lula de Silva, as a proxy for a "political mistrust" variable, and an "aid and commitment" variable designed to capture positive and negative declarations by the IMF concerning possible or agreed support programs to Brazil or Argentina during this period. In the next sections we further justify the use of these variables, while more detailed descriptions are given in the data appendix.

2.4.3 Global Factors: the Developed Market High-Yield Spread

There are at least two reasons why one might expect the developed market high-yield spread—the spread between the return on less-than-investment-grade ("junk") bonds and government or other highly rated bonds—to have an effect on emerging markets. The

[13] Albert and Chib (1993) suggest applying the Gibbs sampler to estimation of Markov-switching models with fixed transition probabilities.

[14] Convergence of the Gibbs sampler was achieved by using 10,000 passes, with the first 1,000 discarded.

first is that movements in the high-yield spread may be a countercyclical leading indicator of economic activity in developed markets, so that a rise in the spread may presage a reduction in economic activity with a concomitant adverse effect on capital flows to emerging markets. The second is that the wedge between the return on high-yield bonds and investment-grade bonds must, more generally, reflect the general attitude towards risk of investors at that point in time—the bigger the wedge, the greater the degree of risk aversion and the less willing will investors be to invest in emerging markets, other things equal.

The idea that the high-yield spread may be a countercyclical leading indicator of economic activity derives from the theory of the financial accelerator.[15] While the finer details of financial accelerator models differ, their central features are reasonably uniform and their key elements may be set out informally as follows: There is some friction present in the financial market, such as asymmetric information or costs of contract enforcement, which, for a wide class of industrial and commercial businesses, introduces a wedge between the cost of external funds and the opportunity cost of internal funds—the "premium for external funds". This premium is an endogenous variable, which depends inversely on the balance-sheet strength of the borrower, since the balance sheet is the key signal through which the creditworthiness of the firm is evaluated. However, balance-sheet strength is itself a positive function of aggregate real economic activity, so that borrowers' financial positions are procyclical and hence movements in the premium for external funds are countercyclical. Thus, as real activity expands, the premium on external funds declines, which, in turn, leads to an amplification of borrower spending, which further accelerates the ex-

[15] See for example Bernanke, Gertler and Gilchrist (1999) and the references therein.

pansion of real activity. This is the basic mechanism of the financial accelerator. In line with the predictions of the theory, Gertler and Lown (1999) and Mody and Taylor (2003) find evidence of a strongly significant and negative relationship between the US high-yield spread and US real activity.

With reference to the Brazilian events under investigation, the credit crunch and the depressed real activity in developed markets associated also with the *Enron* scandal and the related worries about firm accountability, may have generated serious spillovers into developing markets. In particular, the global portfolio rebalancing across classes of investment risk, due to increased risk aversion, pushed up the bond rates of the major Latin American economies. Figure 2.4 shows how the trends of the Brazilian sovereign spread and the and index of developing countries' high-yield spread show quite a similar pattern, especially during the 2002 turmoil. However, the extent of the increase in the developed market high-yield spread seems able to explain only partially the dramatic jump in Brazilian country risk.

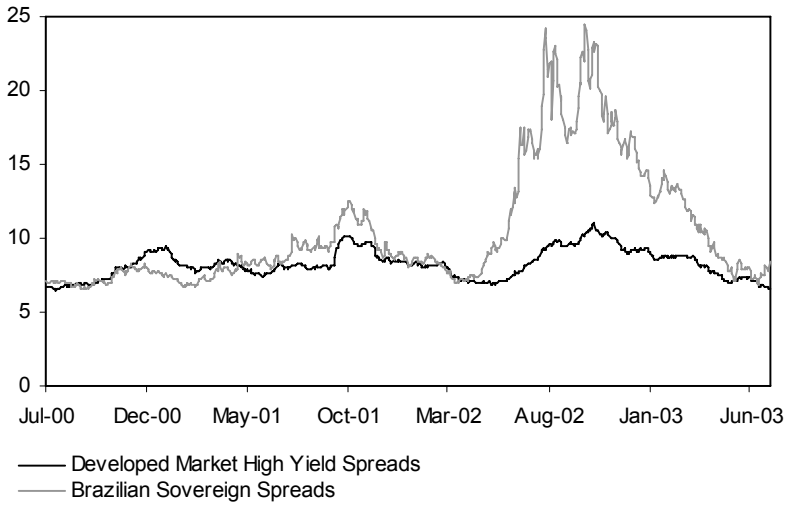


Figure 2.4: Developed-Market High-Yield and Brazilian Sovereign Spreads

Source: JPMorgan

2.4.4 The Role of the Argentine Crisis

Recent studies on financial crises in emerging markets have underlined the existence of substantial spillover effects and contagion episodes amongst countries within a given regional area. This feature can be related to common external macroeconomic shocks, trade and financial linkages among countries or simply a shift in market sentiment. In his analysis of the Mexican crisis, Calvo (1996) supports the idea of herd behaviour in this context, with agents withdrawing their exposure from an entire group of interconnected markets in response to signs of distress in just one of them, rather than bearing the costs of assessing the true state of the underlying fundamentals relating to each market.

This interpretation of contagion episodes could provide a potential explanation for the events in Brazil. At the turn of the century, Argentina, the third largest economy in Latin America and a key trade partner of Brazil, had been in recession since 1998 and was registering a growing public-debt-to-GDP ratio, associated with high political instability, and an overvalued exchange rate with large current account deficits. In January 2002, the country devalued its currency abandoning the parity with the USD introduced under the currency board in 1991. After the country default on public debt, the currency crisis degenerated into a financial one, jeopardised by the asymmetric pesification of bank assets and liabilities with an estimated mismatch of US$ 54 billion. The result of these events was a loss of confidence in the domestic authorities and a corresponding currency and bank run in an attempt to circumvent the introduced restrictions on cash current account withdrawals and the freeze on time deposits.

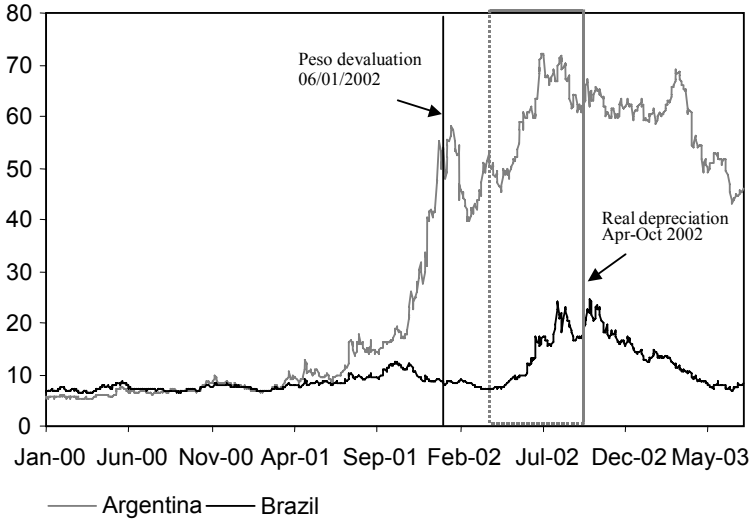


Figure 2.5: Sovereign Spreads in Brazil and Argentina

Source: JPMorgan EMBI Global

Despite the stronger fundamentals of the Brazilian economy, its vast stock of public debt and its vulnerable composition made the country highly exposed to capital flow reversals resulting from changes in market sentiment. The scene was therefore set for a neighbouring crisis such as the Argentine one to generate wide spillovers. Although all emerging markets were experiencing a reduction in foreign investments at the time, due to portfolio reallocation and changes in risk assessment by investors, a "sudden stop" phenomenon impacting upon Argentina and Brazil independently does not seem a plausible explanation. Similarities and closeness probably determined a financial contagion effect of its own which self-reinforced the common external shock. Nevertheless, this view is not shared in academic and institutional circles. According to Krueger (2002), "contagion was

limited because the Argentine default was largely expected. Indeed, the crisis seemed to unfold almost in slow motion. As a result investors had ample opportunity to restructure their portfolios in advance. With the exception of Uruguay, most Latin American banks have maintained only a small exposure to Argentina.". The following empirical analysis may be viewed as an attempt to check the validity of these conclusions, by testing the effect of the Argentine crisis on Brazil, in terms of increased country risk (Figure 2.5)[16]. We therefore included data on Argentine spreads in our empirical analysis.

2.4.5 Electoral Expectations and Country Risk

> "Brazil has implemented strong and consistent macroeconomic policies in recent years that have improved fundamentals [...] Despite these achievements, the uncertain economic environment and some concerns about the course of economic policies following the upcoming presidential elections have put substantial pressure on financial variables" (IMF Press release 02/40, September 2002).

The IMF, as well as a number of other economists, has pointed to the presidential elections held in October 2002 as one of the main sources of economic instability in the country in that period. The drop in net capital flows has in fact been explained as the result, among other reasons, of investors' worries concerning a possible shift in macroeconomic policy following the likely victory of the left-wing candidate, Lula da Silva, in the presidential election contest, mainly due to Lula's past declarations in favour of debt repudiation[17].

Following this view, Razin and Sadka (2002) propose a multiple-equilibria debt-crisis model for Brazil in 2002. The trigger able to coordinate market expectations and induce creditors to change their beliefs about the country's credit worthiness is represented by the

[16] See Pericoli-Sbracia [2001] for a comprehensive review of the relevant literature on contagion.

[17] See Williamson (2002) and (2003) for a wide examination of Brazilian politics in 2002.

forthcoming elections, with a regime change. Once again, the crucial assumption in the model for the occurrence of multiple equilibria is the presence of an indexed debt and the dependence of the country risk on foreign lending. Miller et al. (2003) address the same issue with a model of Bayesian learning, where the voters learn about the candidate's policy preferences and, above all, his attitude towards debt default. These authors allow also for contagion effects from Argentina: the events in the neighbouring country do not cause directly a shift between multiple equilibria; instead, the country debt repudiation raises default expectations in Brazil by shifting prior beliefs about the nature of an incoming left-wing government.

In light of these considerations, we added to our empirical analysis an examination of the effect of a political variable. In particular, we attempted to capture public sentiment towards the likely victory of the socialist candidate and the resulting worries for the country's future fiscal situation by using opinion-poll data in the run-up to the presidential election. The net predominance of Lula's support during the year preceding the elections is clearly described by Figure 2.6, which presents the dynamics of the opinion polls for the favourite candidates in terms of the percentage of those asked who stated that they were going to vote for the candidate is question.

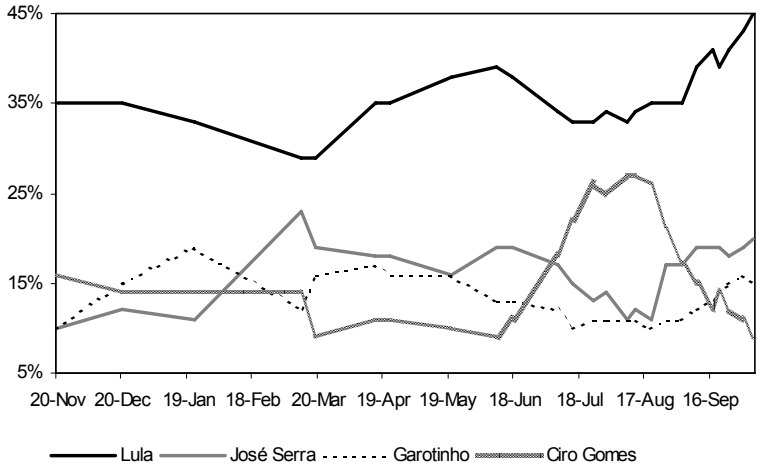


Figure 2.6: Brazilian Political Opinion Polls

Source: IBOPE.

In their paper on the influence of political instability on economic vulnerability, Bussière and Mulder (1999) criticize the use of polls as an indicator because of their lack of credibility, remarking how many election outcomes differ significantly from polls forecasts. Nevertheless, in the context of this study, we believe that polls are the most suitable variable, since we are not looking for exact forecasts of the political outcome, but for the best public signal, available to foreign investors, of the average Brazilian voter preferences for the presidential elections.

2.4.6 What Role for IMF Catalytic Finance?

Finally, we are interested in considering whether the intervention of international financial institutions (IFIs), such as the IMF, can potentially avoid the jump towards the crisis regime or even preclude the existence of multiple equilibria.

In this regard, the literature on currency crises has tended to focus on the role of a common lender of last resort, able to guarantee sovereign debt and honour it in case of country default. This international guarantee has been widely criticised in the aftermath of the Asian crisis, because of the strong moral hazard implications.

A parallel role attributed to IFIs is based on the doctrine of catalytic finance: official assistance to a country in crisis would not only provide the necessary liquidity but would contribute to a strengthening of market sentiment and hence encourage a return of private-sector funding. Recent work by Corsetti et al. (2003) provides an explanation of how such a stabilising mechanism may come into effect, by focusing on the coordination of agents' expectations and government incentives. The main characteristics of the model are the insurgence of a liquidity problem, rather than a solvency one; the effectiveness of IMF support even if the resources available are less than what is needed to close the financing gap; and the possibility of restoring market confidence and thus generating a strong herding effect that is able to solve any coordination failures among creditors. As regards potential distortions in the policymaker behaviour, moral hazard arises only if the true nature of government is misunderstood. In all other cases, a well-intentioned policymaker would find in the IMF's liquidity support, and in the private funding it promotes, the necessary means to realise all the needed reforms. As suggested by Morris and Shin (2003), a useful

integration of the literature on catalytic finance would suggest regarding IMF financial support as conditional on earlier actions and the pre-commitment of the government.

Corsetti et al. (2003) refer to the Brazilian events as a likely example of the effectiveness of IFI intervention in the presence of a country liquidity problem. The successful exit from the currency and financial turmoil of Summer 2002 is likely to have been determined by the positive reaction of the markets to the IMF intervention on September 6, with the approval of a 15-month stand-by credit of about US$31 billion. Furthermore, as has been remarked by the IMF itself, "the commitment that the leading presidential candidates have given to the core elements of the program already appears to have helped market confidence"[18], thus supporting the view that financial aid needs to be linked to a fiscally committed government in order to be effective and credible.

With reference to the Brazilian experience, IMF intervention—coupled with the letters of intent of policymakers and the policy declarations of the presidential candidates—could have had an influence on the vulnerable debt situation as well on the political uncertainty in the pre-election period. The mistrust of Brazilian and international investors regarding the future policies of a left-wing government could have been more than offset by the support action from the Fund.

The commitment not to repudiate the sovereign debt and to undertake all the necessary policy actions to ensure fiscal sustainability, underpinned by the IMF agreement, may have given a strong signal to the public on the true type of government. Chang (2002) shows how financial assistance programs can have a significant impact on domestic poli-

[18] IMF Press release 02/40, September 2002.

tics and, surprisingly, how commitment to a common fiscal policy can especially benefit the pro-labour party, by making candidates more alike. According to this view, the hypothesis of the IMF's intrusion in Brazilian politics, as well as its indirect support to the outgoing administration's candidate was totally unjustified.

From an empirical perspective, these theoretical results offer further possibilities for investigation. As Miller et al. (2003) indicate, "just as bad news from Argentina could increase sovereign spreads [...], so arrangements with the IMF might have the opposite effect". We therefore decided to introduce into our analysis two "aid and commitment" variables, including both Argentine and Brazilian news, in order to test whether positive and negative news concerning IMF intervention, as well as the political commitments from the current government and the presidential candidates, may have had a significant impact on the country default risk.

2.5 Empirical Results

2.5.1 Estimates and Interpretation

Given the previous discussion, the candidate variables to include in our analysis in order to explain the level and shifts in the level of the Brazilian sovereign spread ($BRSPR$) were the Argentine sovereign spread ($ARSPR$), a measure of the perceived probability of Lula's successes in the presidential election, as measured by Brazilian opinion polls ($LULAOPP$), the developed market high-yield spread (HY) and an aid-and-commitment variable which took the value -1 when positive declarations were made concerning IMF

programs to Argentina or Brazil (IMF_YES) and zero otherwise, one which took the value 1 for negative declarations by IMF in this respect—which in the event only concerned negative declarations about Argentina for this period—(IMF_NO)[19]. Further, because of the high level of Argentine spreads over the period in question, we investigated whether there was threshold effect concerning the influence of the Argentine sovereign spread when $ARSPR$ breached a 60% threshold.[20] See the data appendix for further discussion of data sources and methods.

Estimations were carried out using daily data for the period November 20, 2001 to October 28, 2002, with a total sample of 245 observations[21]. With the exceptions of the IMF news dummy variables, all variables were included with a one-period lag in order to preclude any issues of endogeneity of the explanatory variables. In the initial estimations, we included all of these variables in both of the estimated equations (i.e. in both x_t and z_t). After sequentially setting statistically insignificant parameters to zero (using a 5% nominal significance level), we settled on the following preferred specification (where $I(.)$ denotes an indicator variable that takes a value of unity when the indicated inequality is true, and zero otherwise and estimated standard errors are given in parentheses below coefficient estimates):

---

[19] For merely computational reasons, pessimistic news have a positive sign in the dummy variable.

[20] The threshold level of 60% was suggested by estimating a univariate, constant probabilities Markov-switching autoregressive model for $ARSPR$.

[21] We found extremely useful to have access to the Gauss codes by Martin Ellison (European University Institute - April 2000), which we built on in order to develop the program for our model. We thank the author for making them available on line (http://www2.warwick.ac.uk/fac/soc/economics/staff/faculty/ellison/software/tvtp.zip).

$$BRSPR_t = \underset{(0.555)}{6.832} + \underset{(0.301)}{5.687} S_t + \underset{(0.009)}{0.031} I(ARSPR_{t-1} < 60\%)ARSPR_{t-1} +$$

$$+ \underset{(0.004)}{0.125} I(ARSPR_{t-1} \geq 60\%)ARSPR_{t-1} + \varepsilon_t; \tag{6}$$

$$\varepsilon_t \tilde{} N(0, 2.41) \tag{2.7}$$

$$P(S_t = 1) = P(S_t^* \geqslant 0); \tag{2.8}$$

$$S_t^* = \underset{(0.753)}{-5.054} + \underset{(0.178)}{3.288} S_{t-1} + \underset{(0.006)}{0.042} LULAOPP_{t-1} +$$

$$+ \underset{(0.075)}{0.203} HY_{t-1} + \underset{(0.174)}{0.761} IMF\_NO_t + u_t. \tag{2.9}$$

$$u_t \tilde{} N(0, 1) \tag{2.10}$$

The results are extremely encouraging in that all of the estimated coefficients are strongly statistically significantly different from zero and their sign in each case accords with our economic intuition. The coefficient of S_t in equation (6) shows that a switch into the crisis regime entails a jump of about 5.7% in the Brazilian sovereign spread. Interestingly, the only other explanatory variable that was found to be significant in this equation was the Argentine spread. Although this variable does not appear in the final latent variable equation (9), this in fact suggests a strong contagion effect, since a movement of the Argentine

spread above 60% is accompanied by an increase in the slope coefficient by a factor of about four (from 0.031 to 0.125), indicating a contribution to the Brazilian spread in excess of 7%.

With regard to the latent variable equation, the estimated coefficient for the political variable is also strongly significant[22] and shows, as expected, a positive sign, in line with the hypothesis of a direct relationship between Lula's victory chances and the perceived country default risk. As regards the global factor variable, the high-yield spread for developed markets has a strong and significant impact on the probability of being in crisis.

We also find a positive relationship between the probability of being in crisis and the IMF dummy variable for negative news, in favour of the idea that a programme refusal or a pessimistic declaration by the IMF can deeply affect investors' beliefs. In particular, the negative-news variable reports the IMF refusal, on December 5, 2002, to complete the latest review of Argentina's IMF supported program, which would have allowed the country to draw a further US$ 1.3 billion from the IMF[23].

The positive IMF news variable was not found to be statistically significant. This seems to reflect a bias of investors' reaction towards bad news: while reassuring declarations by political candidates and new agreement with the IMF appear to have had no notable impact on market sentiment, negative remarks—which represent sporadic events—can induce strong capital outflows and even trigger a crisis, by decreasing the probability of remaining in a tranquil state. However, a deeper analysis of the estimated probability of

[22] The LULAOPP variable is also economically significant, given it is expressed in percentage.

[23] See IMF Survey 30-23.

being in the crisis state, as drawn in Figure 2.7, shows a temporary tendency of the crisis state to reverse towards the tranquil one in mid-August 2002: this trend precedes closely the announcement of the approval of a $30.4 billion stand-by credit for Brazil and the extension of Argentina's SRF repayment by one year, respectively on September 6 and 5. As a result, we cannot exclude the hypothesis of a leak of information concerning the two IMF interventions in the two weeks immediately before their official announcement to the press.

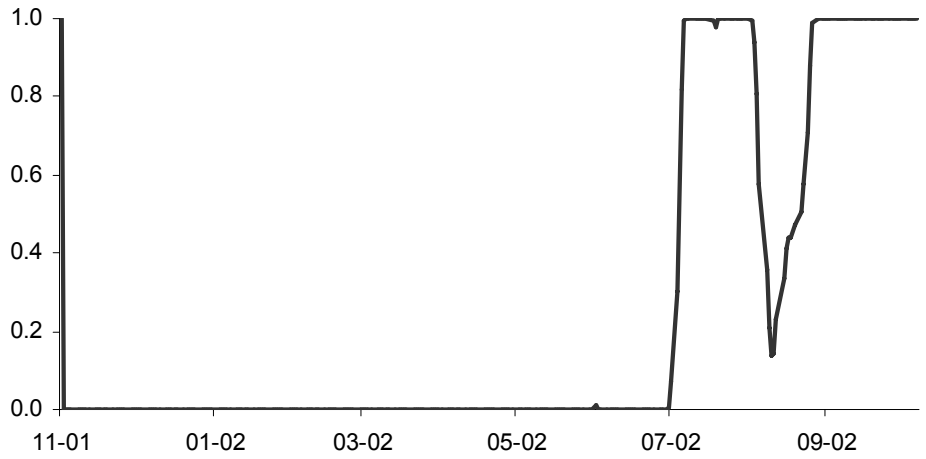


Figure 2.7: Probability of a crisis state

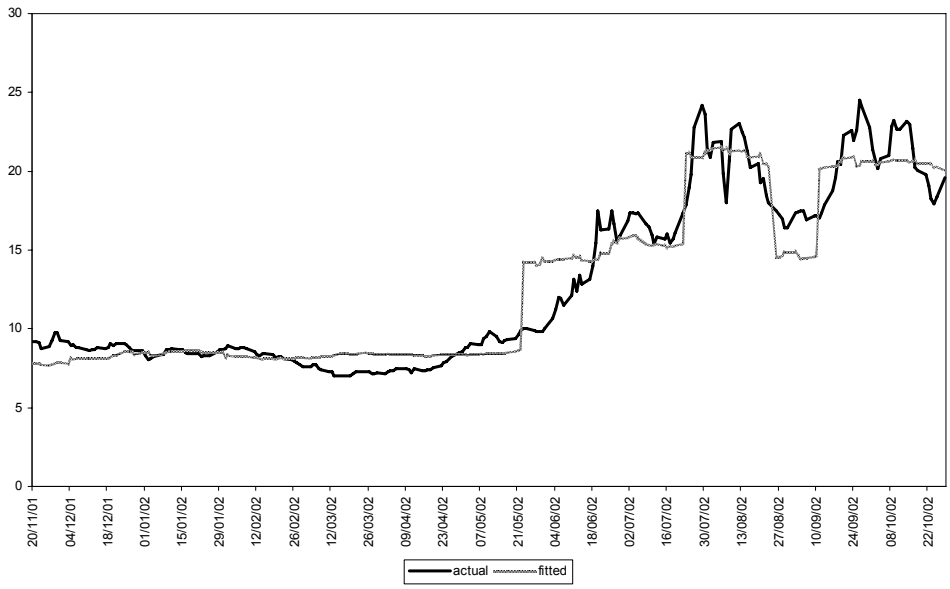


Figure 2.8: Brazilian spreads: actual versus fitted values

The R^2 for the regression is equal to 0.9192, indicating that the selected explanatory variables in the regression account for over 90% of the variation in the actual Brazilian spreads. The good explanatory power of the specification is also illustrated by the plot of the actual versus the fitted values of BRSPR in Figure 2.8.

We have shown how the currency and financial turmoil of 2002 in Brazil was triggered by three kinds of signals: political mistrust by foreign investors concerning the conduct of the expected winner at the presidential elections, the behaviour of high-yield spreads in developed markets and negative news concerning IMF intervention in Argentina. Financial spillovers from Brazil's neighbour, Argentina, appear also to have a role in the 2002 events.

2.6 Concluding remarks

This chapter has investigated the events of 2002 in Brazil, by testing empirically some of the leading interpretations that have been advanced to account for the financial turmoil that characterised that period. Markov switching modelling proved an appropriate econometric tool in the analysis of distress periods. In effect, the possibility of capturing self-fulfilling changes in market behaviour, by means of an endogenous regime-shift selection, allows researchers to test multiple equilibria models empirically. Our estimates provide evidence in favour of financial contagion from the Argentine crisis as well as of political uncertainty during the pre-election period in Brazil. In particular, such instability is interpreted as political mistrust concerning the future conduct of the left-wing leader with respect to the country's fiscal obligations. While political uncertainty contributed to the strong jump in Brazilian bond spreads, the spillover effect from Argentina seems to have affected only the extent of the turmoil in Brazil. The intervention of the IMF, coupled with the fiscal pre-commitments of the domestic authorities and the declarations of the presidential candidates, appeared instead to have a reverting effect on country risk. Nevertheless, while positive news do not show up as a significant factor in avoiding a crisis, negative news and declarations seem to have a deep impact on investors' sentiment and to be determinant in the probability of switching into a crisis regime.

However, this study is by no means exhaustive. Data availability represents the main limitation of the analysis: the use of high-frequency data allows a better representation of the dynamics of investors' behaviour, but reduces the quantity of variables available at the same time. The absence of daily observations for foreign reserves as well as for other fun-

damental variables excludes the possibility of including these variables in the estimations. In addition, even at lower frequency, there is still a difficulty in obtaining country-level data on capital flow movements, while these are widely available for trade statistics. Finally, the chronology of political events and IMF news would need to be compiled with reference to the effective day of disclosure of the information to the markets, in order to take account of the possible existence of "insider trading" and information leaks, given the large numbers of people and institutions involved in the decision process.

Finally, the results of our study suggest a number of policy considerations. The significant role of contagion between South American economies clearly needs to be taken constantly into account by IFIs, notably the IMF, in their intervention strategies, perhaps by means of more effective policy coordination in the region. Moreover, the potential for IMF declarations and agreements to act as a common public signal able to coordinate market expectations and hence generate a catalytic-finance effect strongly encourages further research on this topic.

2.A Appendix: Data Definitions and Sources

2.A.1 Sovereign spreads

The data source for bond spreads is the JP Morgan Chase's Emerging Markets Global Bond Index (EMBI Global). According to JP Morgan's methodology brief[24], the EMBI Global includes U.S.-dollar-denominated Brady bonds, Eurobonds, traded loans and local market debt instruments issued by sovereign and quasi-sovereign entities. The weight of each individual issue in the country-by-country index is determined on the basis of market capitalisation. The use of a bond-spread index rather than the difference between an individual sovereign bond yield and the US Treasury bill yield closest to it in issue date and maturity allows for a better homogeneity of the data sample for different countries over time .The EMBI Global Spreads are available at daily frequency since January 1, 1998 for Argentina and Brazil (These two variables are referred to in the analysis as ARSPR and BRSPR).

2.A.2 Developed Market high-yield Spread

The data source for this variable is the JP Morgan Chase's Developed Market high-yield Summary Spread to Worst. The JPMorgan Developed Market HY Index represents all US$ denominated corporate issues, with issuers domiciled in countries labelled as industrial by JPMorgan's Economic Research Group. The spread is given by the difference between the high-yield and yield of the Treasuries. Issues included in the index must be rated "5B"

[24] J.P.Morgan (1999)

or lower. That is, the highest Moody's/S&P ratings are Baa1/BB+ or Ba1/BBB+. The inception date of the index is January 1994.

2.A.3 Political variable

Data on opinion polls during the 2002 presidential electoral campaign in Brazil have been obtained from IBOPE (*Instituto Brasileiro de Opinião Pública e Estatística*). Daily observations have been linearly interpolated from the data series available, from November 20, 2001 to October 28, 2002, with a total sample of 245 observations (The variable is named LULAOPP in the analysis). In particular, the following list of *Pesquisas De Opinião Pública Sobre Assuntos Políticos/ Administrativos* (Public Opinion Surveys on Political/Administrative Affairs) has been considered:

OPP169 19/nov/2001; OPP203 19/dez/2001; OPP006 21/jan/2002; OPP035 11/mar/2002; OPP051 18/mar/2002; OPP079 21/apr/2002; OPP085 21/apr/2002; OPP107 19/mai/2002; OPP135 19/jun/2002; OPP172 07/jul/2002; OPP184 14/jul/2002; OPP202 23/jul/2002; OPP218 30/jul/2002; OPP225 08/ago/2002; OPP235 13/ago/2002; OPP249 20/ago/2002; OPP279 27/ago/2002; OPP305 03/set/2002; OPP329 09/set/2002; OPP351 17/set/2002; OPP357 20/set/2002; OPP385 24/set/2002; OPP450 01/out/2002; OPP422 05/out/2002; OPP438 6/out/2002; Result 1st round; OPP469 15/out/2002; OPP513 22/out/2002; OPP524 26/out/2002; OPP537 27/out/2002; Result 2nd round.

2.A.4 Aid and commitment variable

In the analysis we consider two different "aid and commitment" variables. The first dummy variable (IMF_YES) takes a value equal to minus one each time the IMF released optimistic declarations regarding possible or agreed supported programs to Brazil or Argentina. The variable also includes the issue dates of letter of intents by the Brazilian national authorities. In addition, news concerning the disbursement delay of the IMF to Argentina on December 5, 2001 and any related event has been included in an another dummy variable (IMF_NO), with a value of one. The main news sources were the IMF website and major Brazilian newspapers and news and information agencies. For the purposes of the analysis and according to the other variable sample-periods, we used observations from between November 20 2001 through October 28 2002, for a total of twenty news items for Brazil and eleven for Argentina.

The events captured by the two variables are listed below by country and typology of news.

POSITIVE NEWS:

Brazil

19 November 2001 Statement by Mr. Pedro Malan, Minister of Finance of Brazil, at the International Monetary and Financial Committee, Ottawa, November 17, 2001

30 November 2001 Brazil Letter of Intent, Memorandum of Economic Policies, and Technical Memorandum of Understanding

23 January 2002 News Brief: IMF Completes First Review of Stand-by Arrangement with Brazil

28 January 2002 IMF Survey: IMF approves review of Brazil's performance

07 February 2002 Public Information Notice: IMF Concludes 2001 Article IV Consultation with Brazil

04 March 2002 Brazil Letter of Intent, Memorandum of Economic Policies, and Technical Memorandum of Understanding, March 4, 2002

26 March 2002 News Brief: IMF Completes Second Review of Stand-By Arrangement with Brazil

22 April 2002 Statement by Mr. Pedro Malan, Minister of Finance of Brazil to the International Monetary and Financial Committee, Washington, D. C., April, 20, 2002

27 May 2002 IMF Survey: Brazil positioned to speed up growth

05 June 2002 Brazil – Letter of Intent, Memorandum of Economic and Financial Policies, Technical Memorandum of Understanding

14 June 2002 Brazil – Letter of Intent

18 June 2002 News Brief: IMF Completes Third Review of Stand-By Arrangement with Brazil

21 June 2002 Lula's "Carta ao Povo Brasileiro"

24 June 2002 IMF Survey: Brazil can draw $10 billion;

23 July 2002 Lula's "Compromisso com a soberania o emprego e a seguranca do povo brasileiro"

07 August 2002 News Brief: IMF Managing Director Köhler Confirms Agreement with Brazil

29 August 2002 Brazil – Letter of Intent, Memorandum of Economic Policies, and Technical Memorandum of Understanding

06 September 2002 Press Release: IMF Approves US$30.4 Billion Stand-By Credit for Brazil

16 September 2002 IMF Survey: Brazil loan

30 September 2002 IMFC Statement by Mr. Pedro Malan, Minister of Finance of Brazil, on behalf of the Constituency comprising Brazil, Colombia, Dominican Republic, Ecuador, Guyana, Haiti, Panama, Suriname, and Trinidad and Tobago

28 October 2002 News Brief: IMF Managing Director Köhler Congratulates Brazil's President-Elect

Argentina

16 January 2002 Press Release: IMF Extends Argentina's SRF Repayment by One Year

28 January 2002 IMF Survey: IMF extends Argentine debt deadline

08 February 2002 News Brief: IMF's Köhler Welcomes Remes Visit, says IMF Working Closely with Argentina

13 February 2002 News Brief: IMF's Köhler: Good Start to New Relationship with Argentina

04 March 2002 News Brief: IMF Sending Mission to Argentina
11 March 2002 IMF Survey: IMF mission to Argentina
15 March 2002 News Brief: Press Statement by the IMF Mission to Argentina
25 March 2002 IMF Survey: IMF statement on Argentina
10 April 2002 IMF Encourages Argentine Reforms, By Thomas C. Dawson, Director, Ext. Rel. Dept, IMF; Introductory Remarks on the Role of the IMF Mission in Argentina by Anoop Singh, Dir. for Special Operations, IMF
17 April 2002 Statement by the IMF Mission to Argentina
21 May 2002 News Brief: IMF Managing Director meets Argentine Economy Minister, Board extends repayment of SRF
28 June 2002 News Brief: IMF's Köhler Welcomes Progress in Talks with Argentina
08 July 2002 IMF Survey: Köhler on Argentina
10 July 2002 News Brief: IMF Managing Director Horst Köhler Announces Advisory Group on Argentina
15 July 2002 Press Release: IMF Extends Argentina's SRF Repayment by One Year
22 July 2002 IMF Survey: Advisory panel for Argentina
29 July 2002 News Brief: IMF Managing Director Köhler Welcomes Independent Advisors' Report on Argentina
05 August 2002 IMF Survey: Krueger on Argentina
05 September 2002 Press Release: IMF Extends Argentina's SRF Repayment by One Year

NEGATIVE NEWS:

Argentina

05 December 2001 Transcript of a Press Briefing by Thomas C. Dawson, Director, External Relations Department, IMF
06 December 2001 Transcript of a Press Briefing by Thomas C. Dawson, Director, External Relations Department, IMF
10 December 2001 IMF Survey: Argentine disbursement delayed

Chapter 3
Are There Thresholds of Current Account Adjustment in the G7?

3.1 Introduction

The sustainability and adjustment of current account imbalances among the world's major industrialized countries is a subject that is receiving considerable attention among policy-makers, financial market practitioners and academics. At more than $600 billion and nearly 6 percent of US GDP, the US current account deficit attracts the most focus, but there are also material current account imbalances in other deficit countries such as the UK and in surplus countries such as Japan and Germany.

Some respected experts have expressed concern that current account imbalances of this magnitude and persistence indicate that the global economy is operating in a "danger zone" in which disruptive and volatile reactions in currency, bond, and equity markets are likely to result. For example, in C. Fred Bergsten (2002, p.5) has argued that "research at both the Federal Reserve Board and the Institute for International Economics reveals that industrial countries, including the United States, enter a "danger zone" of current account unsustainability when their deficits reach 4-5 percent of GDP[...] At these levels, corrective forces tend to arise either spontaneously from market forces or by policy action". Other observers have made a similar point, arguing that there is a "threshold" current account imbalance beyond which current account adjustment must ultimately take place, even if

evidence of adjustment is scarce or non-existent before the threshold is reached. This point of view is represented clearly in a recent survey paper on this subject prepared by the Fed Reserve Bank of Kansas City (Holman (2001, p.16)):

> While there is considerable uncertainty about the precise threshold [...] a current account deficit greater than 4.2 percent of GDP is unsustainable. This estimate, based on the 1980s and early 1990s, represents the average threshold at which current account deficits in several industrialized economies started to narrow after trending up for a sustained period.

Existing empirical work on this subject is suggestive but is not in fact specifically aimed at answering the question "are there thresholds of current account adjustment", or exploring its implications. Influential papers by Milesi Ferretti-Razin (1998) and Freund (2000) employ a careful and informative methodology to pull together a set of empirical regularities about how adjustments of large current account deficits have taken place in previous "episodes" which meet certain ex ante criteria. For example, in order for a current account deficit adjustment episode (called a reversal) to be included in the Freund sample, it must meet the following four criteria:

1. The current account deficit exceeded two percent of GDP before the reversal.
2. The average deficit was reduced by at least two percent of GDP over three years (from the minimum to the three-year average).
3. The maximum deficit in the five years after the reversal was not larger than the minimum deficit in the three years before the reversal.
4. The current account was reduced by at least one third.

These are very similar to the criteria introduced by Milesi-Ferretti and Razin (1998) in their study. Their motivation for focusing on the adjustment of large current account deficits that meet these criteria is explained as follows (p. 12):

> In the definition of reversal events we want to capture large and persistent improvements in the current account imbalance, that go beyond short-run current account fluctuations as a result of consumption smoothing. The underlying idea is that "large" events provide more information on determinants of reductions in current account deficits than short run fluctuations.

The work of Milesi Ferretti - Razin, Freund, and – using a somewhat different methodology Mann (2002) – has had an impact on the way that policymakers discuss current account adjustment, especially in the context of the record US deficits recorded in recent years. For example, Federal Reserve Chairman Greenspan (2003), citing Freund's work has said:

> [W]hat do we know about whether the process of reining in our current account deficit will be benign to the economies of the United States and the world? According to a Federal Reserve staff study, current account deficits that emerged among developed countries since 1980 have risen as high as double-digit percentages of GDP before markets enforced a reversal. The median high has been about 5 percent of GDP.

While much can be and has been learned by studying past episodes of adjustment of large current account deficits (as defined by the criteria used by Milesi Ferretti-Razin and Freund), there remains a number of unresolved empirical questions pertaining to the modeling, estimation, and interpretation of the current adjustment process among the large industrialized countries. These questions include:

- Does the process of adjusting to current account deficits differ from the process of adjusting to current account surpluses? (*does sign matter?*)

- Does the process of adjusting to ”large” current account imbalances differ from the process of adjusting to smaller current imbalances? (*does size matter?*)
- If so, is there a way to estimate how large is ”large” and does this estimate differ from country to country? (*does one size fit all?*)
- Is the absence of evidence about the adjustment of a large current account imbalance evidence in favor of the sustainability of said large imbalance? (*is the absence of evidence evidence of sustainability?*)

It is the aim of this chapter to provide an empirical framework that can be used to begin to answer questions such as these. We will argue that, for any particular country, all four of these issues are in fact intrinsically related to one another and to the specification of the econometric model which best describes that country's current account dynamics. If the current account, suitably scaled by net output (GDP net of investment and government purchases), is a linear, stationary stochastic process with a constant unconditional mean, as is often assumed in empirical work, then the answers to these four questions are straightforward: 'no', 'no', 'moot', and 'yes'.

An immediate implication of stationarity is that any current account/net output ratio not equal to the unconditional mean is unsustainable by the definition of a stationary stochastic process. This applies to surpluses as well as deficits. However, as an empirical matter, the dynamic process by which the current account adjusts to its unconditional mean depends crucially on whether the process is linear or nonlinear. In particular, if the process is linear, adjustment is symmetric above and below the long-run equilibrium, and the speed

of adjustment is independent of the magnitude of the displacement from long-run equilibrium (the unconditional mean). For a linear, stationary current account/net output process, there is nothing to be gained by just focusing on the adjustment of current account deficits and excluding the data on adjustment to surpluses (all relative to the unconditional mean current account/net output ratio which may be either positive or negative). Moreover, there is no reason to focus on the adjustment to "large" deficits as providing different or more information than episodes of adjustment to small deficits (relative to the unconditional mean) since all episodes provide the same information. Finally, as should be obvious by now, for a linear stationary stochastic process there is no particular "threshold" beyond which markets and/or shifts in policy force a reversal and below which adjustment is absent.

By contrast, if the stationary stochastic process which governs the current account adjustment to its long mean is non-linear, then both the "sign" and "size" of the current account imbalance *does* matter for the adjustment process, and the size of the current account imbalance beyond which adjustment takes place may well be country specific (as alluded to by Chairman Greenspan and as is suggested by the empirical work cited above). Finally, if the stationary stochastic process is non-linear, absence of evidence of adjustment of a large current account imbalance is not evidence of the absence of the ultimate adjustment of the imbalance.

There is a tractable and testable nonlinear time series model that conveniently exhibits all of the features of the current account adjustment process that have been the focus of recent discussions, and that nests as a special case the linear stationary stochastic process model for the current account that is often assumed in empirical work. It is the thresh-

old auto regression model introduced in Tong (1978) and studied extensively by Hansen (1996,1999a, 1999b). For a stationary stochastic threshold model with mean μ and thresholds $\underline{\delta}$ and $\overline{\delta}$, there is no tendency for $ca = current\ account/net\ output - \mu$ to adjust to its mean of 0 unless it has crossed either the threshold $\underline{\delta}$ or the threshold $\overline{\delta}$. In the 'regime' with $\underline{\delta} < ca < \overline{\delta}$, deficits or surpluses (relative to μ) persist, and there is no tendency for imbalances to revert. However, the absence of evidence of mean reversion in this regime is not evidence that deficits or surpluses relative to μ are 'sustainable' since, by stationarity, the only sustainable current account imbalance is equal to the unconditional mean.

In a threshold model, a necessary condition for adjustment to commence is for ca to cross either the deficit threshold $\underline{\delta}$ or the surplus threshold $\overline{\delta}$, parameters which can be estimated from the data, not imposed ex ante. In the deficit adjustment regime, $ca < \underline{\delta}$, and $ca_t = \underline{\rho} ca_{t-1} + \varepsilon_t$. Adjustment continues until ca reaches $\underline{\delta}$ at which point any further adjustment is driven by shocks to ε_t. In the surplus regime adjustment regime, $ca > \overline{\delta}$, and $ca_t = \overline{\rho} ca_{t-1} + \varepsilon_t$. Adjustment continues until ca reaches $\overline{\delta}$ at which point any further adjustment is driven by shocks to ε_t. Evidently, in a threshold model, the sign and size of the ca imbalance can matter, thresholds can differ across countries, and the absence of evidence of adjustment is not the evidence of absence of future adjustment of the ca imbalance.

The plan of the chapter is as follows. In Section 2 we review some basic empirical predictions of the modern workhorse model of the current account, the rational expectations, intertemporal approach model developed in Sachs (1981, 1982), estimated by Sheffrin and Woo (1990), and recently extended by Kano (2003). The basic prediction of this

model, once one allows for permanent shocks to the level of net output as in Campbell and Deaton (1989), is that the ratio of the current account to net output (GDP less investment less government purchases) should be a stationary stochastic process with an unconditional mean determined by the relationship between the real interest rate and the per capita rate of growth. We also argue that a general equilibrium, two-country version of the Weil (1989) infinite horizon, overlapping generations model of the current account – a model in which the global real interest rate and the net foreign asset or liability position of each country is endogenously determined – also has the prediction that the current account to net output ratio is constant in steady state and determined by underlying parameters such as rates of time preference, the steady state rate of global growth, and the relative size of the two countries. In this chapter, we will follow most of the empirical work in this area and take the stationarity of the current account to net output ratio as given. The question at the heart of the present chapter is whether or not the stationary stochastic process which describes the current account to net output ratio in the G7 countries features linear or nonlinear adjustment to the unconditional mean. We conclude Section 2 by presenting, for each G7 country the results of a non-parametric statistical test of the null hypothesis of a linear adjustment of the current account to net output ratio against the alternative of non-linear adjustment using quarterly data for the sample 1979:1 - 2003:3. This is an application of a test for nonlinearity developed by Terasvirta (1994). For the G7 countries in our sample, we find statistically significant evidence against the null of linear adjustment of the current account to net output ratio and in favor of the alternative of nonlinear adjustment.

In Section 3 of the chapter, we estimate for each G7 country a threshold autoregressive model of the current account to net output ratio, allowing for country-specific thresholds of current account surplus and deficit adjustment in each country (as suggested, for example, by Chairman Greenspan's comments), and also allowing for country specific means for the ratio of the current account to net output (as suggested, for example, by the general equilibrium version of the Weil model reviewed in Section 2). Our main findings in this section are as follows. For most of the G7 countries, we find significant evidence of threshold effects in current account adjustment. We also find that we cannot reject the null hypothesis of a random walk for the current account imbalance in each country when that ratio does not exceed (in absolute value) the country specific surplus and deficit thresholds (relative to the country specific mean) estimated for that country. For most of the G7 countries, unless the current account imbalance is 'too large' – as suggested by Milesi-Ferretti and Razin (1998) – there does not appear to be a systematic tendency for adjustment to occur. A further advantage of our approach is that we can estimate from the data how large a current imbalance has to be before this imbalance triggers an adjustment, and we can allow these estimated thresholds to differ across countries. In fact we find substantial cross country variation in the surplus and deficit thresholds that trigger current account adjustment in each country. We also find evidence of cross country and cross regime variation in the autoregressive dynamics estimated during adjustment regimes for each country

In Section 4, we investigate what happens to the probability distributions of nominal exchange rate changes, stock price index changes, and long term interest rate differentials during the various current account adjustment regimes that we estimate for each country

in Section 3. The motivation is to determine whether or not crossing the current account adjustment threshold is itself associated with shifts in the probability distributions for exchange rates, stock prices, and interest differentials. We specifically account for – and allow for current account regime-specific shifts in – autoregressive conditional heteroskedasticity as well as for shifts in the mean by estimating generalized autoregressive conditionally heteroskedastic (GARCH) models for nominal exchange rate changes, stock prices changes, and interest differentials. We also in this section explore, for the US, whether or not the expectation of a future adjustment in the current account imbalance is associated with a present shift in the probability distribution of exchange rates, stock prices, or interest differentials. We proxy this by including in the GARCH models two dummy variables (one for deficits and one for surpluses) which represent the distance between the current account imbalance and its country-specific mean when the imbalance is between the thresholds.

In Section 5, we draw on our empirical results to take stock of the present US current account deficit. Our empirical results indicate that compared to other G7 countries, the US over our sample exhibited relatively wide thresholds within which current account adjustment is absent and relatively slow speeds of adjustment once these thresholds, especially the deficit threshold, are crossed. Moreover, the present US current account deficit substantially exceeds – and has for some time – our estimated thresholds of current account deficit adjustment for the US. We explore several possible explanations. The first is that the threshold model, while a useful description of current account adjustment for other G7 countries, does not apply to the US and that the present deficit of nearly 6 percent of GDP is in fact sustainable. The second explanation is that there are thresholds of current account

adjustment for the US, but that adjustment has been delayed over the past several years, due to unusual circumstances that were not in evidence during the sample over which the models were estimated, 1979-2003. These circumstances could include: (i) the low level of global real interest rates (which support higher levels of investment and lower levels of saving in the US than would be the case with historically average or above average real interest rates); (ii) the more muted and less uniform decline in the dollar than occurred, for example, during the 1985 - 1987 Plaza-Louvre episode (reflecting the intervention activities of Asian central banks); (iii) the fact that the US continues to run a substantial surplus in dividends, interest, and profits on its stock of foreign assets compared with the dividends, interest, and profits that it pays out on its much larger stock of foreign liabilities; (iv) the adjustment in the net foreign liability position of the US that occurs as a result of dollar depreciation (which in 2003 offset almost 80 percent of that year's current account deficit). We review and evaluate these potential explanations for the absence of adjustment to date in the US current account deficit even though it has passed well beyond thresholds that would have triggered adjustments in other G7 countries.

Section 6 provides some concluding remarks.

3.2 A Test for Non-Linear Current Account Adjustment

3.2.1 Theoretical Considerations

In our empirical work, we shall be modeling the dynamics of G7 current account adjustment. However, it is important to take a stand as to exactly what it is to which G7 current

account imbalances are adjusting. In this chapter, we draw on the implications for long-run current account equilibrium of the workhorse intertemporal model of the current account (Sachs, 1981; Sheffrin and Woo, 1990, via Campbell, 1987). This model can be written $CA_t = -E_t \sum(1+r)^{-i}\Delta Z_{t+i}$ where $Z_t = Y_t - I_t - G_t$ is the level of net output. The intertemporal approach models have been estimated and tested many times, and their high frequency implications – that current account dynamics are fully described by the discounted sum of future changes in net output – are usually rejected. However, we argue that the intertemporal model, properly specified to allow for stationarity in long run growth rates, contains an important insight about the long run behavior of the current account. It would seem preferable to model $\Delta \log Z_t = \Delta z_t$ as stationary. Following Campbell and Deaton (1989), it is straightforward to show (Kano, 2003) that the log-linear approximation of the intertemporal approach model is given by $CA_t/Z_t \approx -E_t \sum(1+r-g)^{-i}\Delta z_{t+i}$ where g is the unconditional mean of Δz_t. Note that if the log difference of net output is stationary, it is the current account to net output *ratio* which is stationary, not simply the current account itself. This seems like a more sensible long-run equilibrium condition than to assume that the current account itself is stationary.

The intertemporal approach model is partial equilibrium and is usually studied for the special case in which r is equal to the rate of time preference. However, the basic prediction of that model – that the ratio CA/Z is constant in the long run – also holds in the steady state of a two-country version of Weil's (1989) infinite horizon overlapping generations model. As shown in Obstfeld and Rogoff (1994, page 188), the Weil model with discount factor β implies that the steady state current account to net output ratio is constant and

given by $CA/Z = (n+g)\theta$ where n is the rate of population growth, g is the rate of net output growth, and θ is the endogenous ratio of net foreign assets to net output given by the solution to $\theta[1-(1+r)\beta/(1+n)(1+g)] = [(1+r)\beta-(1+g)]/(1+n)(1+g)(r-g)$. Now imagine two such economies trading goods and bonds with one another that differ in two respects: size and the discount factor. Let $\beta_1 < \beta_2$ and suppose that country 2 is larger than country 1. It is easy to show that in the steady state of a two-country version of the Weil model, the β_1 smaller country will run a steady state current account to net output deficit and the larger more patient β_2 country will run a steady state current account to net output surplus. Based on these considerations, we shall assume that for each G7 country, the ratio CA/Z is stationary and allow for country specific means in the CA/Z ratio.

3.2.2 Testing for Non-Linearities in G7 Current Account Adjustment

This chapter is an empirical study of G7 current account adjustment, based on quarterly data for the period 1979:1 to 2003:3 (the data available when we began our study in the fall of 2003). We choose our starting date to begin six years after the advent of floating exchange rates and the initial globalization of the international capital market that occurred at that time and in conjunction with the first oil shock. The data in the analysis are obtained from the International Financial Statistics Database by the IMF. All variables are seasonally adjusted and expressed in national currency. According to national account statistics, the current account variable is estimated as the sum of net exports and net primary income from abroad (NPIA); net output is obtained by subtracting Government consumption expenditure and gross fixed capital formation (investment) to GDP.

We test for non-linearity in G7 current account/net output adjustment following the non-parametric test for nonlinearity developed by Luukkonen et al. (1988) and Terasvirta (1994). These authors propose a Lagrange Multiplier test for a third-order Taylor approximation to the regression function of the form: $ca_t = \beta_{00} + \sum_{i=1}^{p}[\beta_{1i}ca_{t-1} + \beta_{2i}ca_{t-i}ca_{t-d} + \beta_{3i}ca_{t-i}ca_{t-d}^2 + \beta_{4i}ca_{t-i}ca_{t-d}^3] + \varepsilon_t$. This artificial regression allows to identify general nonlinearity through the significance of the higher-order terms. The main advantage of this type of test is that it can be carried out by simple OLS and that – despite being designed for smooth transition regressions – is sensitive to a wide range of non-linearities (Granger and Terasvirta, 1993), although there is reason to suspect that the power of the test may be weak against some nonlinear alternatives. The results of this test are reported in Table 3.1.

Country	**Terasvirta Linearity Tests:**
CAN	Marginal significance level 0.369
FRA	Marginal significance level 0.029
GER	Marginal significance level 0.035
ITA	Marginal significance level 0.136
JAP	Marginal significance level 0.027
UK	Marginal significance level 0.184
US	Marginal significance level 0.069

Table 3.1: Terasvirta Linearity Tests

Hence, evidence of nonlinear adjustment is indicated at the 5% significance level for France, Germany and Japan, and at the 7% level for the US.

Using the multivariate bootstrap test procedure developed by Hansen (1997), the null hypothesis of linear adjustment in all countries is rejected at the 14% level. Given the possibly poor power characteristics of these tests, therefore, we felt encouraged to investigate the estimation of nonlinear models more directly.

3.3 Estimating and Testing Thresholds Models of G7 Current Account Adjustment

In this section of the chapter, we estimate and test for each G7 country a threshold autoregression model of the current-account-to-net-output ratio using the univariate approach developed in Hansen (1996). We allow for and estimate country-specific means, country and regime-specific thresholds, and country and regime specific dynamic adjustment once the current account has crossed either of the thresholds. Letting $ca = CA/Z - \mu$, we write the equilibrium threshold autoregressive (TAR) model as

$$\begin{aligned} ca_t &= \overline{\rho} \times 1\{ca_{t-d}, \overline{\delta}\} \times ca_{t-1} + \underline{\rho} \times 1\{ca_{t-d}, \underline{\delta}\} \times ca_{t-1} + \\ &\quad +(1 - 1\{ca_{t-d}, \overline{\delta}\}) \times (1 - 1\{ca_{t-d}, \underline{\delta}\}) \times ca_{t-1} + e_t \end{aligned} \tag{3.11}$$

where $1\{ca_{t-d}, \overline{\delta}\}$ is an indicator function that takes on a value of 1 when $ca_{t-d} > \overline{\delta} \geqslant 0$ (and zero otherwise) and $1\{ca_{t-d}, \underline{\delta}\}$ is an indicator function that takes on a value of 1 when $ca_{t-d} < \underline{\delta} \leqslant 0$ (and zero otherwise). This approach postulates that the persistence

of the current account imbalance in a country may depend upon whether or not the current account imbalance has crossed a surplus 'threshold' of $\overline{\delta} \geqslant 0$ or a deficit threshold of $\underline{\delta} \leqslant 0$. We note that a special case of the threshold model is the case in which $\overline{\delta} = \underline{\delta} = 0$ and $\overline{\rho} = \underline{\rho} < 1$ in which case it collapses to a linear stationary AR(1) process. We experimented with a threshold TAR(2) specification but found in general the second lag terms to be insignificant, and thus confine our presentation to the TAR(1) models. We also select a delay parameter d of two quarters as this maximises the fit of the regression in each case.

The threshold model can potentially identify three regimes of current account adjustment: a surplus adjustment regime, a deficit adjustment regime, and an 'inertia' regime $\overline{\delta} < \mathbf{ca}_{t-2} < \underline{\delta}$ in which the current account appears to follow a random walk. In a more general smooth threshold transition autoregressive or STAR model (e.g. Taylor, Peel and Sarno, 2001), the speed of adjustment does not increase discontinuously at the threshold; rather, the further way is the current-account-to-GDP ratio from its long-run mean, the faster the current account imbalance adjusts. Interestingly, when we experimented with estimating smooth transmission models, we found they did not capture G7 current account dynamics in a sensible way. As we shall report next, there does in fact appear to be important, discrete threshold effects which influence current account adjustment.

Before presenting the results, we will discuss some issues involved in the estimation and testing of these model for a system comprised of the G7 countries. The ca variables for the G7 group are first demeaned, in order to allow for the existence of long-run deficit/surplus means for each country rather than a zero ca balance. A non-zero mean proves to be applicable for all G7 countries, with the single exception of Italy. In partic-

ular, we detect a structural break in the German series in 1991, corresponding with the German unification and the resulting change in the country national accounts; we account for the break by allowing two different means in the current account for the pre- and post-unification periods.

The two asymmetric thresholds in the TAR model are selected jointly by minimisation of the overall sum of squared errors. The estimation method involves a double grid search over ca. Following Hansen (1997), the range for the grid search is selected a priori to contain ca observations in between the 15th ($\underline{ca}$) and the 85th percentile ($\overline{ca}$). This reduction in the grid range is needed in order to avoid sorting too few observations in one regime for extreme values of the thresholds. As a result, the appropriate ranges are defined as $\overline{\mathbf{R}} = [\mu, \overline{ca}]$ and $\underline{\mathbf{R}} = [\mu, \underline{ca}]$, for $\overline{\delta}$ and $\underline{\delta}$ respectively.

As the minimisation process for a three-regime/ two-threshold TAR process is numerically intensive, we rely on the estimation methodology proposed by Hansen (1999a) for multiple thresholds. This consists of a three-stage grid search, where the second-stage estimation of the two-threshold model is made conditional on the first-stage single-threshold estimate of δ (either $\overline{\delta}$ or $\underline{\delta}$), the third stage being used as a refinement.

Furthermore, final estimates of slope parameters and standard errors for the G7 group of countries are obtained by seemingly unrelated regression (SUR) estimation, in order to allow for potential correlation between the disturbances of the different ca equations, due to common unobservable factors.

Once the thresholds have been selected, according to standard asymptotic theory, (1) is linear in the parameters. As with any simple dummy-variable regression, it can be

estimated by linear methods. However, statistical inference in a TAR model bears the difficulty that the thresholds $\overline{\delta}$ and $\underline{\delta}$ may be not identified under the null hypothesis in question (Davies, 1987). In this case, the usual χ^2 distribution needs to be replaced by an approximated empirical distribution obtained by bootstrapping the residuals (Hansen 1997). In particular, artificial observations are calibrated using the restricted estimates and are then used to obtained new estimates of the restricted and unrestricted model (for an application, see Peel and Taylor 2002). The percentage of bootstrap samples – we run 1000 replications – for which the simulated likelihood-ratio statistics exceeds the actual one forms the bootstrap approximation to the p-value of the test statistic under question.

The estimation and testing results are presented in Table 3.2. First the test results: when we test the null hypothesis a single threshold for all countries versus the alternative hypothesis of two thresholds, we reject the null hypothesis in favor of the alternative. This is consistent with three regimes for each country - a surplus adjustment regime, a deficit adjustment regime, and an inertia (absence of adjustment) regime. Second, when we test the hypothesis that the current account follows a random walk inside the 'inertia' regime against the alternative that it follows a mean reverting autoregressive process inside the inertia regime (a more general formulation of the threshold model) we are unable to reject the null of a random walk inside the inertia regime. In summary, the statistical tests find evidence of non-linear current account adjustment and also identify significant thresholds beyond which current account adjustment takes place.

Threshold models of de-meaned CA/NO

CA/NO Q1 1979- Q3 2003	Thresholds (asymmetric band)		Slope coefficients (estimation by SUR)			Means	
COUNTRY	Upper threshold	Lower threshold	above	band	below	'Surplus'	'Deficit'
CANADA	**1.41**	**-4.05**	0.927 (0.048)	1.000	0.930 (0.060)		**-1.792**
FRANCE	**2.13**	**-1.13**	0.931 (0.048)	1.000	0.910 (0.045)	**1.646**	
GERMANY	**2.84**	**0.00**	0.880 (0.070)	1.000	0.827 (0.064)	**6.185** Pre-1991 **1.496** Post-1991	
ITALY	**0.00**	**-0.37**	0.944 (0.058)	1.000	0.867 (0.059)		**-0.269**
JAPAN	**0.84**	**-0.18**	0.908 (0.058)	1.000	0.894 (0.037)	**3.951**	
UK	**1.08**	**0.00**	0.777 (0.073)	1.000	0.929 (0.064)		**-1.764**
US	**2.15**	**-2.18**	0.907 (0.039)	1.000	0.973 (0.034)		**-2.011**

SE in brackets

Bootstrap:

LR-test for band coefficient equal to 1 (SUR): marg.signif. level = 0.520
LR-test for single threshold (SUR): marg. signif. level = 0.004

Table 3.2: Thresholds Models of Current Account

We now discuss the parameter estimates for the threshold models estimated for each G7 country. To repeat, these estimates allow for country-specific means, country and regime-specific thresholds, and country and regime specific autoregressive dynamics. A number of interesting results are obtained. First, as suggested by Chairman Greenspan's comment cited above, we see there is wide cross-country variation in the estimated current account deficit adjustment thresholds. For example, the estimated deficit adjustment threshold for the US is -2.18 percent of net output, while for Japan it is only -0.18 percent of net out-

put. This means that empirically, there is no evidence from these estimates of systematic adjustment in the US current account deficit until the deficit exceeds -4.19 percent of net output (equal to the mean of -2.01 plus the threshold of -2.18), while for Japan, adjustment begins to take place when the surplus falls below 3.77 percent of net output (equal to the mean of 3.95 plus the deficit threshold of -0.18). We estimate a similar pattern for the other 'structural' surplus countries, France and Germany. For France, we estimate that adjustment begins to take place once the surplus falls below 0.51 percent of net output; for Germany adjustment begins to take place once the surplus falls below the mean of 6.19 before unification and 1.19 percent after unification. Second, we see that for most G7 countries, there are thresholds of adjustment to current account surpluses as well as for current account deficits. Third, we see from Table 3.2 substantial cross-country variation in the estimated autoregressive dynamics once countries cross their current account deficit or surplus thresholds. For deficit adjustment episodes, the estimated autoregressive coefficients range from 0.827 for Germany to 0.973 for the US. For surplus adjustment episodes, the estimated autoregressive coefficients range from 0.777 in the UK to 0.944 in Italy.

HALF LIFE OF DISPLACEMENT FROM DEFICIT THRESHOLD (IN QUARTERS)	1 percent	2 percent	3 percent
Canada	1.14	2.49	3.30
France	2.84	4.08	4.79
Germany	3.65	3.64	3.64
Italy	3.18	3.84	4.13
Japan	4.79	5.48	5.69
UK	9.41	9.41	9.41
G6 Avg	4.17	4.82	5.16
Us	6.25	9.99	12.49
HALF LIFE OF DISPLACEMENT FROM SURPLUS THRESHOLD (IN QUARTERS)			
Canada	3.07	4.58	5.48
France	2.43	3.88	4.84
Germany	1.09	1.81	2.32
Italy	12.03	12.03	12.03
Japan	3.29	4.50	5.13
UK	1.09	1.56	1.82
G6 Avg	3.83	4.72	5.27
Us	1.77	2.82	3.53

Table 3.3: Half Life of Displacement

In the top panel of Table 3.3, we compute the half life of 1 , 2, and 3 percent of net output displacements of the current account imbalance from the deficit threshold. In our equilibrium threshold model the speed of adjustment to a given displacement from the deficit (or surplus) threshold is a function of the distance between the imbalances and the unconditional mean, not just to the threshold itself (as for example would be the case for a so called band threshold model). As is evident from the table, the US stands out in terms of the slow speed of adjustment to current account deficits, even when it is adjusting. For example, in response to a 2 percent of GDP displacement of the US current account from the estimated deficit threshold of -2.18 percent (to a deficit of -4.18 percent of net output), it takes the US nearly 10 quarters on average to close 1 percentage point of that displacement, whereas for the average G6 country (G7 minus US), it takes fewer than 5 quarters to close

such a displacement. In the bottom panel of Table 3.3, we compute the half life of 1, 2, and 3 percent of net output displacements of the current account imbalance from the upper (surplus) threshold. As before, we estimate substantial cross-country variation in the speeds of adjustment to displacements of the current account away from the adjustment thresholds. Note that the US actually adjusts faster than the G6 average to current account surpluses.

PERCENT OF SAMPLE SPENT IN EACH REGIME

	CANADA	FRANCE	GERMANY	ITALY	JAPAN	UK	G6 AVG	US
SURPLUS	34	23	20	51	36	37	34	20
INERTIA	48	35	20	3	30	17	25	63
DEFICIT	18	42	60	46	34	46	41	17

ADJUSTMENT PER QUARTER DURING ADJUSTMENT REGIMES
(Measured from peak and as percent of net output)

	CANADA	FRANCE	GERMANY	ITALY	JAPAN	UK	G6 AVG	US
SURPLUS	0.687	0.507	1.081	0.467	0.336	0.644	0.620333	0.303
DEFICIT	0.604	0.246	0.693	0.575	0.361	0.612	0.515167	0.327

Table 3.4: Summary Statistics

In Table 3.4, we present some summary statistics for the three current account regimes estimated for each G7 country. We see that the average G6 (excluding the US) country spent only roughly 25 percent of the 1979- 2003 sample in the inertia regime and thus spent 75 percent of the sample adjusting to either current account surpluses (34 percent of the sample) or deficits (41 percent of the sample). Of course, there is cross-country variation, but the G6 country spending the maximum time in the inertia regime was Canada, which spent 48 percent of sample in the inertia regime. The US, by contrast, spent a full 63 percent of the sample in the inertia regime, and only 17 percent of the sample adjusting to current account deficits, and 20 percent of the time adjusting to current account surpluses. The

bottom panel of Table 3.4 reports, for each country, the average adjustment per quarter that actually occurred during the sample (as a percentage of net output) when that country was estimated to be in a deficit adjustment regime or a surplus adjustment regime. These adjustments are measured from the peak current account imbalance reached during the adjustment episode to the level reached when the adjustment regime concludes. Thus, for the average G6 country, once current account deficits (relative to mean) peak and begin to contract, they adjust at an average rate of 0.51 percent of net output per quarter (2 percent of net output per year) until adjustment concludes with the current account imbalance crossing the deficit adjustment threshold. The table also shows that for the G6, on average, once current account surpluses peak and begin to contract, they adjust at an even faster average rate 0.62 percent of net output per quarter (2.4 percent of net output per year) until adjustment concludes with the current account imbalance crossing the surplus adjustment threshold. Evidently, adjustment of current account imbalances in the US data is much more sluggish than the G6 average, with the US current account imbalance falling by roughly 0.3 percent of net output during each quarter (1.2 percent per year) that the US is in an adjustment regime.

To summarize the results of this Section, having tested and found evidence of nonlinearity in G7 current account adjustment data, we estimated for each G7 country a threshold autoregressive model which allows for asymmetric, country-specific thresholds, country specific means, and regime and country specific speeds of adjustment. We find evidence in favor of deficit as well as surplus thresholds for most countries, as well as evidence of substantial cross-country differences in the amount of time spent in the three different

regimes, as well as in the pace at which adjustments occur. Compared with other G7 countries, the US has large thresholds of current account adjustment, spends relatively little time in adjustment regimes, and adjusts slowly even when in those imbalance adjustment regimes. In the next section of the chapter, we explore what happens to the probability distributions of exchange rates, stock prices, and interest rate differentials during current account adjustment regimes in each country.

3.4 Exchange Rates, Stock Prices, and Interest Rates During Current Account Adjustment Regimes

In this section, we investigate what happens to the probability distributions of nominal exchange rate changes, stock price index changes, and long term interest rate differentials during the various current account adjustment regimes that we estimate for each country in Section 3. The motivation is to determine whether or not crossing the current account adjustment threshold is itself associated with shifts in the probability distributions for exchange rates, stock prices, and interest differentials. We specifically account for – and allow for current account regime specific shifts in – autoregressive conditional heteroskedasticity as well as for shifts in the mean by estimating GARCH models for nominal exchange rate changes, stock prices changes, and interest differentials. We also in this section explore, for the US, whether or not the expectation of a *future* adjustment in the current account imbalance is associated with a present shift in the probability distribution for exchange rates, stock prices, or interest differentials.

Switching models of exchange rates were introduced in Engel and Hamilton (1990). They hypothesized that the log difference in the nominal exchange rate is a stochastic process with a regime-specific mean and a regime specific (but constant) variance. In their model, the regimes themselves are unobservable states; the probability that the exchange rate is in a particular regime is inferred from the exchange rate data itself. Our approach is different, but similarly motivated. Having found evidence of three regimes of current account adjustment for each G7 country, we estimate and test whether or not being in a current account adjustment regime is associated with shifts in the drift and variance of exchange rate changes for that country. We allow for autoregressive conditional heteroskedasticity in exchange rate changes. We estimate similar models for the log difference in stock price changes and for long term interest rate differentials, allowing for regime specific drifts and variances.

The GARCH models we estimate in this section are of the form

$$\begin{aligned} \Delta_t &= d + d1DUMS_t + d2DUMD_t + u_t \qquad (3.12) \\ \sigma_t^2 &= c + au_{t-1}^2 + b\sigma_{t-1}^2 + c1DUMS_t + c2DUMD_t \end{aligned}$$

where $DUMD_t$ is a dummy variable that takes on a value of 1 when a country is in a deficit adjustment regime, $DUMS_t$ is a dummy variable that takes on a value of 1 when a country is in a surplus adjustment regime, σ_t^2 is the conditional variance of u_t, and Δ_t is the log difference in the exchange rate, the log difference in the equity price index, or the interest rate differential (adjusted for first order autocorrelation) observed at a monthly frequency. Thus, in each quarter in which a country is in a particular regime,

there will be three observations on the monthly change in the asset price during that quarter. Because Italy and France were part of the EMS during most of the sample, the behavior of their exchange rates and interest rates reflected their EMS commitments to stabilize their exchange rates vis-a-vis Germany. We exclude them from the analysis of this section. Estimation is by maximum likelihood. For each country, we report the results for the log (change) in the trade weighted exchange rate, the (log change) in a broad stock market index, and the differential between each county's long term interest rate and G7 average (adjusted for first order autocorrelation). When significant, we also report the results for key bilateral exchange rates. In what follows '*' indicates significance at the 5 percent level, '**' significance at the 10 percent level, and '***' at the 15 percent level. Data sources are the IFS for long-term interest rates and Bloomberg for exchange rates and stock market indices. The sample is monthly from 1979:2 to 2003:9 with some exceptions as noted below.

3.4.1 Results

US RESULTS

For the US dollar index, we see that the estimated coefficient on the surplus regime dummy is positive and the estimated coefficient on the deficit regime dummy is negative (Figure 3.1). This means that the dollar index tends to appreciate during US surplus adjustment regimes, and to depreciate during US deficit adjustment regimes, although the coefficients are not measured precisely. For the pound, we estimate a statistically significant shift in the probability distribution of exchange rate changes that coincides with US

surplus adjustment regimes, in favor of an appreciation of the dollar relative to the pound. For the Canadian dollar, we estimate a statistically significant shift in the probability distribution of exchange rate changes that coincides with US deficit adjustment regimes, in favor of a depreciation of the dollar relative to the Canadian dollar. We also estimate a statistically significant rise in the volatility of the Canadian dollar exchange rate that coincides with US deficit adjustment regimes. For US equity prices, we estimate a significant (at the 12 percent level) fall in equity returns during US current account deficit adjustment regimes. We also estimate a significant rise in equity volatility that occurs during US current account adjustment regimes. For long term interest rate differentials, we do estimate a significant increase in volatility during US current account surplus adjustment regimes.

Japanese Results

For the Yen index, we see that the estimated coefficient on the Japan current account surplus adjustment regime dummy is positive and significant, indicating that the Yen index tends to appreciate during Japan's current account surplus adjustment regimes (Figure 3.2). For the Dollar-Yen exchange rate, we estimate a statistically significant increase in exchange rate volatility during both Japan surplus adjustment regimes and Japan deficit adjustment regimes. We also obtain point estimates that suggest that the yen tends to appreciate relative to the dollar during Japanese current account surplus regimes and to depreciate during Japanese current account deficit adjustment regimes, although these coefficients are not measured precisely. For Japanese equity prices, we estimate a significant fall in equity volatility during Japan current account deficit adjustment regimes. For long term interest rate differentials, we do estimate a significant increase in volatility during both Japan's cur-

rent account surplus adjustment regimes and current account deficit adjustment regimes. We also estimate a significant widening in Japanese long term interest differential (it becomes larger in absolute value) during Japan's current account surplus adjustment regimes, as well as a widening during Japan's current account deficit adjustment regimes (although the latter is not significant).

German Results

For the volatility of the DM index through 1998:12, we see that the estimated coefficient on the German current account deficit adjustment regime dummy is positive and significant (Figure 3.3). For the Dollar-DM exchange rate estimated through 1998:12, we estimate a statistically significant depreciation of the DM during German current account deficit adjustment regimes. For German equity prices, we estimate a significant fall in equity volatility during German current account deficit adjustment regimes. For long- term interest rate differentials, we do estimate a significant increase in volatility during German current account deficit adjustment regimes. German interest rate differentials increase in absolute value during deficit adjustment regimes in before unification, and narrow after unification. We split the sample at unification because of an obvious shift in the mean of the interest differential series at that time.

UK and Canadian Results

For the Canadian dollar index, we see that the estimated coefficient on the Canadian current account deficit adjustment regime dummy is negative and significant, indicating that the CAD index tends to depreciate during Canada's current account deficit adjustment

regimes (Figure 3.4). For the US Dollar-Canada exchange rate, we estimate a similar result but it is not statistically significant. For the UK, the most noteworthy result is a significant increase in equity returns during current account surplus adjustment regimes, a fall in equity volatility during UK current account surplus adjustment regimes, and a rise in equity volatility during UK current account deficit adjustment regimes (Figure 3.5). Because of a break in the UK equity price data series at 1984:1, the UK equity sample is 1984:1 - 2003:9.

Summary of Results

In this subsection, we have reported evidence of statistically significant shifts in the mean and variance of the probability distribution of several G7 exchange rates, equity prices, and interest rate differentials that occur in conjunction the current account adjustment regimes estimated in section 3. Our approach cannot answer the question of which triggers what, but we do find evidence that regimes of current account adjustment do coincide with shifts in the distribution of some important asset prices. The estimates that are significant tend to show exchange rate depreciation during current account deficit regimes and exchange rate appreciation during current account surplus regimes. We also find statistically significant increases in exchange rate volatility during current account deficit adjustment regimes for the US, Japan, and Germany. For equity markets, we estimate that current account deficit adjustment regimes are associated with significantly lower US equity returns and higher US equity volatility, while in the UK, equity returns are higher during current account surplus adjust regimes, equity volatility is lower, while UK equity volatility is higher during current account deficit adjustment regimes.

Asset Prices During US Current Account Adjustment Regimes

US Dollar Index

$$\Delta_t = -.0004 + \underset{(.0028)}{.0035DUMS_t} - \underset{(.0025)}{.0028DUMD_t} + u_t$$

$$\sigma^2_t = .0001 - .0325\, u^2_{t-1} + .5976\sigma^2_{t-1} + \underset{(.00003)}{.00002DUMS_t} - \underset{(.00003)}{.00002\, DUMD_t}$$

Pound per Dollar

$$\Delta_t = -.0013 + \underset{(.0044)*}{.0101DUMS_t} - \underset{(.0038)}{.0019DUMD_t} + u_t$$

$$\sigma^2_t = .0002 + .2151\, u^2_{t-1} + .6013\sigma^2_{t-1} + \underset{(.0001)}{.0001DUMS_t} - \underset{(.00007)}{.00002\, DUMD_t}$$

Canadian Dollars per US Dollar

$$\Delta_t = .0009 + \underset{(.0019)}{.0006DUMS_t} - \underset{(.0025)**}{.0044DUMD_t} + u_t$$

$$\sigma^2_t = .0002 - .0161\, u^2_{t-1} - .5754\sigma^2_{t-1} + \underset{(.0001)}{.00001DUMS_t} + \underset{(.00007)*}{.0002\, DUMD_t}$$

Equity Prices

$$\Delta_t = .0107 - \underset{(.0061)}{.0029DUMS_t} - \underset{(.0091)***}{.0139DUMD_t} + u_t$$

$$\sigma^2_t = .0014 + .0004\, u^2_{t-1} + .0681\sigma^2_{t-1} + \underset{(.0004)}{.00027DUMS_t} + \underset{(.0011)*}{.00223\, DUMD_t}$$

Long Term Interest Differentials

$$\Delta_t = .0094 - \underset{(.0304)}{.0154DUMS_t} - \underset{(.0181)}{.0014DUMD_t} + u_t$$

$$\sigma^2_t = .0002 - .0177\, u^2_{t-1} + .9788\sigma^2_{t-1} + \underset{(.0009)*}{.00305DUMS_t} + \underset{(.00014)}{.00007\, DUMD_t}$$

Figure 3.1: Asset Prices During US Current Account Adjustment

Asset Prices During Japan Current Account Adjustment Regimes

Yen Index

$$\Delta_t = -.0016 + \underset{(.0034)^*}{.0093DUMS_t} + \underset{(.0031)}{.0005DUMD_t} + u_t$$

$$\sigma^2_t = .0006 - .2115\,u^2_{t-1} - .2848\sigma^2_{t-1} + \underset{(.00013)}{.00012DUMS_t} - \underset{(.00012)}{.00005\,DUMD_t}$$

Dollar per Yen

$$\Delta_t = .0008 + \underset{(.0050)}{.0066DUMS_t} - \underset{(.0048)}{.0044DUMD_t} + u_t$$

$$\sigma^2_t = .00001 - .0095\,u^2_{t-1} + .9383\sigma^2_{t-1} + \underset{(.00005)^*}{.00012DUMS_t} + \underset{(.00003)^*}{.00008\,DUMD_t}$$

Equity Prices

$$\Delta_t = -.0031 + \underset{(.0084)}{.0105DUMS_t} + \underset{(.0076)}{.0093DUMD_t} + u_t$$

$$\sigma^2_t = .0006 + .1245\,u^2_{t-1} + .7605\sigma^2_{t-1} - \underset{(.0003)}{.00017DUMS_t} - \underset{(.00029)^{***}}{.00044\,DUMD_t}$$

Long Term Interest Differentials

$$\Delta_t = -.1045 - \underset{(.0344)}{.0153DUMS_t} - \underset{(.0371)^*}{.0844DUMD_t} + u_t$$

$$\sigma^2_t = .0049 + .0082\,u^2_{t-1} - .1245\sigma^2_{t-1} + \underset{(.0142)^*}{.028796DUMS_t} + \underset{(.01493)^*}{.03240\,DUMD_t}$$

Figure 3.2: Asset Prices During Japan Current Account Adjustment

Asset Prices During German Current Account Adjustment Regimes

DM Index

$$\Delta_t = .0021 - \underset{(.0014)}{.0013DUMS_t} - \underset{(.0012)}{.0012DUMD_t} + u_t$$

$$\sigma^2_t = .00002 + .0886\, u^2_{t-1} + .1619\sigma^2_{t-1} + \underset{(.00001)}{.00001DUMS_t} + \underset{(.00001)^*}{.00003\, DUMD_t}$$

Dollar per DM

$$\Delta_t = -.0058 - \underset{(.0066)}{.0013DUMS_t} - \underset{(.0053)^{***}}{.0082DUMD_t} + u_t$$

$$\sigma^2_t = .00127 + .0921\, u^2_{t-1} - .2801\sigma^2_{t-1} - \underset{(.0004)}{.00004DUMS_t} + \underset{(.00031)}{.00008\, DUMD_t}$$

Equity Prices

$$\Delta_t = .0037 - \underset{(.0144)}{.0025DUMS_t} + \underset{(.0102)}{.0053DUMD_t} + u_t$$

$$\sigma^2_t = .0015 + .0726\, u^2_{t-1} + .7386\sigma^2_{t-1} - \underset{(.0006)}{.00026DUMS_t} - \underset{(.00051)^*}{.00115\, DUMD_t}$$

Long Term Interest Differentials
1979:1 – 1990:12

$$\Delta_t = -.0129 - \underset{(.0481)}{.0282DUMS_t} - \underset{(.0541)^*}{.2147DUMD_t} + u_t$$

$$\sigma^2_t = .0242 + .2351\, u^2_{t-1} - .0644\sigma^2_{t-1} + \underset{(.0122)}{.01303DUMS_t} + \underset{(.02499)^{***}}{.03635\, DUMD_t}$$

1991:1 – 1998:12

$$\Delta_t = .0074 - \underset{(.0927)}{.0619DUMS_t} - \underset{(.0247)^{***}}{.0358DUMD_t} + u_t$$

$$\sigma^2_t = -.0001 + .0804\, u^2_{t-1} + .7183\sigma^2_{t-1} + \underset{(.0152)}{.01583DUMS_t} + \underset{(.00294)^{***}}{.00455\, DUMD_t}$$

Figure 3.3: Asset Prices During German Current Account Adjustment

Asset Prices During UK Current Account Adjustment Regimes

Pound Index

$$\Delta_t = \underset{}{-.0013} + \underset{(.0029)}{.0012DUMS_t} + \underset{(.0028)}{.0019DUMD_t} + u_t$$

$$\sigma^2_t = .00011 + .2775\,u^2_{t-1} + .5646\sigma^2_{t-1} - \underset{(.00005)***}{.00007DUMS_t} - \underset{(.00005)**}{.00008\,DUMD_t}$$

Dollar per Pound

$$\Delta_t = .0049 - \underset{(.0044)*}{.0093DUMS_t} - \underset{(.0045)}{.0035DUMD_t} + u_t$$

$$\sigma^2_t = .00024 + .1959\,u^2_{t-1} + .5747\sigma^2_{t-1} - \underset{(.0001)}{.00004DUMS_t} + \underset{(.0001)}{.00001\,DUMD_t}$$

Equity Prices

$$\Delta_t = -.0006 + \underset{(.0082)*}{.0185DUMS_t} + \underset{(.0081)}{.0048DUMD_t} + u_t$$

$$\sigma^2_t = .0040 + .0224\,u^2_{t-1} - .8964\sigma^2_{t-1} - \underset{(.0003)*}{.00084DUMS_t} + \underset{(.00070)}{.00091\,DUMD_t}$$

Long Term Interest Differentials

$$\Delta_t = .0312 + \underset{(.032)}{.0073DUMS_t} + \underset{(.028)}{.0177DUMD_t} + u_t$$

$$\sigma^2_t = .00037 + .0461\,u^2_{t-1} + .9402\sigma^2_{t-1} + \underset{(.0018)}{.00048DUMS_t} - \underset{(.0012)}{.00037\,DUMD_t}$$

Figure 3.4: Asset Prices During UK Current Account Adjustment

Asset Prices During Canada Current Account Adjustment Regimes

CAD Index

$$\Delta_t = .0002 - \underset{(.0014)}{.0015DUMS_t} - \underset{(.0017)***}{.0025DUMD_t} + u_t$$

$$\sigma^2_t = .00004 + .1961\, u^2_{t-1} + .4708\sigma^2_{t-1} - \underset{(.00001)}{.000002DUMS_t} + \underset{(.00002)}{.000002\, DUMD_t}$$

US Dollar per Canadian Dollar

$$\Delta_t = .0003 - \underset{(.0014)}{.0018DUMS_t} - \underset{(.0018)}{.0021DUMD_t} + u_t$$

$$\sigma^2_t = .00001 + .0608\, u^2_{t-1} + .8727\sigma^2_{t-1} + \underset{(.00006)}{.00004DUMS_t} + \underset{(.00005)}{.00002\, DUMD_t}$$

Equity Prices

$$\Delta_t = .0051 + \underset{(.0067)}{.0030DUMS_t} - \underset{(.0065)}{.0030DUMD_t} + u_t$$

$$\sigma^2_t = .0007 + .0534\, u^2_{t-1} + .7576\sigma^2_{t-1} - \underset{(.0002)***}{.00041DUMS_t} - \underset{(.00047)}{.00062\, DUMD_t}$$

Long Term Interest Differentials

$$\Delta_t = .1855 - \underset{(.0605)}{.0429DUMS_t} + \underset{(.0331)}{.0300DUMD_t} + u_t$$

$$\sigma^2_t = .0124 + .1002\, u^2_{t-1} + .6336\sigma^2_{t-1} + \underset{(.0033)***}{.05082DUMS_t} + \underset{(.00396)}{.00013\, DUMD_t}$$

Figure 3.5: Asset Prices During Canada Current Account Adjustment

3.4.2 Do Expectations of Future US Current Account Adjustment Trigger Adjustment in Present Asset Prices?

We now explore, for the US, whether or not the expectation of a *future* adjustment in the current account imbalance is associated with a *present* shift in the probability distribution for exchange rates, stock prices, or interest differentials. As discussed previously, compared with other G7 countries, the US has wide thresholds of current account adjustment, spends relatively little time in adjustment regimes, and – as shown in Table 3.4 – adjusts slowly even when in deficit or surplus adjustment regimes. To capture the hypothesis that expectations of future current account adjustment may have an impact on present asset prices, we augment our basic GARCH specification to include two additional dummy variables. Let $DUMBD$ equal one when $-2.18 < ca < -1$ and let $DUMBS$ equal one when $1 < ca < 2.15$. Thus $DUMBD$ equals one when the current account deficit is more than one percentage point below its mean but still less (in absolute value) than the deficit threshold, while $DUMBD$ equals one when the current account is more than one percentage point above its mean but still less (in absolute value) than the surplus threshold. Our specification becomes

$$\begin{aligned} \Delta_t &= d + d1DUMS_t + d2DUMD_t + d3DUMBS_t + d4DUMBD_t + u_t \qquad (3.13) \\ \sigma_t^2 &= c + au_{t-1}^2 + b\sigma_{t-1}^2 + c1DUMS_t + c2DUMD_t + c3DUMBS_t + c4DUMBD_t \end{aligned}$$

In order to focus on significant results, we proceed in two steps. In the first step, we estimate specification (3). In the second step, we drop any dummy variable that in the first

stage estimate is not significant at the 15 percent level or better. The results are reported in Figure 3.6.

Asset Prices Before and During US Current Account Adjustment Regimes

US Dollar Index

$$\Delta_t = .0006 - \underset{(.0033)^{**}}{.0064}DUMBD_t + u_t$$

$$\sigma^2_t = .00012 - .05\,u^2_{t-1} + .7083\sigma^2_{t-1} - \underset{(.00003)^{*}}{.00006}DUMBS_t$$

Equity Prices

$$\Delta_t = .0115 - \underset{(.0087)^{***}}{.0131}DUMD_t + u_t$$

$$\sigma^2_t = .0015 + .0058\,u^2_{t-1} + .1106\sigma^2_{t-1} - \underset{(.0003)^{*}}{.0007}DUMBS_t + \underset{(.00097)^{*}}{.0019}\,DUMD_t$$

Long Term Interest Differentials

$$\Delta_t = -.0020 + \underset{(.0194)^{*}}{.0384}DUMBS_t$$

$$\sigma^2_t = .0003 + .0241\,u^2_{t-1} + .9418\sigma^2_{t-1}$$

Figure 3.6: Asset Prices Before and During US Current Account Adjustment Regimes

From Figure 3.6, we see that when current account deficits are large but *before* the US enters a current account deficit adjustment regime, the dollar index starts to depreciate, at a pace of roughly 7 percent per year. We also see that the volatility of the dollar index is lower when deficits are small but before the US enters a current account surplus adjustment regime. As for equity prices, the results reported in Figure 3.1 are robust to the inclusion of the two additional dummy variables. We continue to find a significant negative effect of current account deficit adjustment regimes on equity returns, and a significant positive

effect on equity volatility. Interestingly, we also find that equity volatility is lower when deficits are small but before they have entered a current account surplus adjustment regime. Finally, we see that long-term interest differentials in favor of the US are larger when current account deficits are small.

3.5 Assessing the Present US Current Account Deficit

In this section we draw on our empirical results to take stock of the present US current account deficit. Our empirical results indicate that compared to other G7 countries, the US over our sample exhibited relatively wide thresholds within which current account adjustment is absent and relatively slow speeds of adjustment once these thresholds, especially the deficit threshold, are crossed. Moreover, the present US current account deficit substantially exceeds – and has for some time – our estimated thresholds of current account deficit adjustment for the US. We explore several possible explanations. The first is that the threshold model, while a useful description of current account adjustment for other G7 countries, does not apply to the US and that the present deficit of nearly 6 percent of GDP is in fact sustainable. The second explanation is that there are thresholds of current account adjustment for the US, but that adjustment has been delayed over the past several years, due to unusual circumstances that were not in evidence during the sample over which the models were estimated, 1979-2003. These circumstances could include: (i) the low level of global real interest rates (which support higher levels of investment and lower levels of saving in the US than would be the case with historically average or above average real interest rates); (ii) the more muted and less uniform decline in the dollar than occurred, for

example during the 1985 - 1987 Plaza-Louvre episode (reflecting the intervention activities of Asian central banks); (iii) the fact that the US continues to run a substantial surplus in dividends, interest, and profits on its stock of foreign assets compared with the dividends, interest, and profits that is pays out on its much larger stock of foreign liabilities; (iv) the adjustment in the net foreign liability position of the US that occurs as a result of dollar depreciation (which in 2003 offset almost 80 percent of that years current account deficit). We review and evaluate these potential explanations for the absence of adjustment to date in the US current account deficit even though it has passed well beyond the thresholds that would have triggered adjustments in other G7 countries. We begin by reviewing the data on the US net foreign liability position.

Almost all claims held by foreigners against the US are dollar denominated, while US claims against the rest of the world are denominated in foreign currency. Thus, as has been emphasized by Pierre Olivier Gourinchas and Helene Rey, a real depreciation of the dollar, by increasing the real value of US holdings of foreign assets relative to foreign holdings of US assets (which of course are dollar denominated liabilities of the US) is an important channel of international adjustment, over and above the impact of said real depreciation on the trade balance. This channel operates by narrowing the gap between the market value of foreign claims against the US and the market value of US claims against the rest of the world. In effect, because of the willingness on the part of the rest of the world to lend to the US in the form of dollar denominated debt and equity instruments, there is a transfer of wealth to the US from the rest of the world as result of a real depreciation of the dollar, all other things – including other asset prices – equal, a qualification to

which we return below. It is important to note that while the US benefits from this 'transfer' effect which increases the real value of US assets relative to US liabilities, there is of course another implication of real dollar depreciation which is the terms of trade deterioration that results from it. This terms of trade deterioration lowers the real purchasing power of any given flow of US income, and it increases the relative price of imported inputs to US based production. In addition, as Obstfeld and Rogoff have emphasized, moving toward current account sustainability requires that resources be shifted from non tradable to tradable production. Empirically, this channel of international adjustment is potentially quite important in complementing the traditional channel in which the factors that contribute to a narrowing of the current account deficit also result in a real depreciation of the dollar.

Every year, the US Commerce Department reports data on the net foreign liability position of the US, and it provides detail on the revaluation of US assets and liabilities that occurs as a result of exchange rate movements, as well as asset price changes. The data on net foreign assets and liabilities is subject to substantial revisions. However, until quite recently – April 2005 – the Commerce Department did not go back and revise the exchange rate and asset price revaluation attributions to make them consistent with the revised data on foreign assets and liabilities. However, the Commerce Department has now revised the exchange rate and asset price revaluation attributions to make them consistent with the revised data on foreign assets and liabilities. The newly released data are reported in Table 3.5 and they tell an interesting story.

Components of Change in the Net International Investment Position, With Direct Investment at Market Value, 1989-2004

Year	Position Beginning	Changes in position					Position Ending
		Attributable to				Total	
		Financial flows	Valuation adjustments				
			Price changes	Exchange-rate changes[1]	Other changes[2]		
		(a)	(b)	(c)	(d)	(a+b+c+d)	
1988	....	....	....	....	....	....	10,466
1989	10,466	-49,545	7,129	-15,392	355	-57,453	-46,987
1990	-46,987	-60,337	-148,620	57,042	34,407	-117,508	-164,495
1991	-164,495	-46,421	-95,789	4,643	41,243	-96,324	-260,819
1992	-260,819	-96,253	-75,554	-74,991	55,312	-191,486	-452,305
1993	-452,305	-81,489	292,716	-21,969	118,779	308,037	-144,268
1994	-144,268	-127,052	23,172	73,069	39,828	9,017	-135,251
1995	-135,251	-86,298	-152,461	39,018	29,156	-170,585	-305,836
1996	-305,836	-137,687	84,188	-66,076	65,387	-54,188	-360,024
1997	-360,024	-221,334	-92,069	-207,625	58,320	-462,708	-822,732
1998[r]	-822,732	-69,740	-287,874	68,120	41,457	-248,037	-1,070,769
1999[r]	-1,070,769	-236,148	329,872	-125,970	65,778	33,332	-1,037,437
2000[r]	-1,037,437	-486,373	133,716	-270,594	79,681	-543,570	-1,581,007
2001[r]	-1,581,007	-400,243	-224,184	-151,685	17,671	-758,441	-2,339,448
2002[r]	-2,339,448	-500,316	-59,582	231,247	212,985	-115,666	-2,455,114
2003[r]	-2,455,114	-560,846	-1,716	415,507	229,599	82,744	-2,372,370
2004[p]	-2,372,370	-584,597	146,514	272,278	-4,070	-169,875	-2,542,245

[p] Preliminary.
[r] Revised.
1. Represents gains or losses on foreign-currency-denominated assets and liabilities due to their revaluation at current exchange rates.
2. Includes changes in coverage, capital gains and losses of direct investment affiliates, and other adjustments to the value of assets and liabilities.

Table 3.5: Components of Change in Net International Investment Position

We begin with the most recent data available as of the time of writing, for year end 2003 (data for end 2004 are preliminary). The US began 2003 with gross foreign assets of $6.6 trillion and gross foreign liabilities of $9.2, for a stock of net foreign liabilities of $2.6 trillion. During that year the US ran a current account deficit of $530 billion which, after adjustment for errors and omissions, resulted in a net capital inflow of $560 billion. In a simple textbook model which abstracts from asset price or exchange rate changes, this should have resulted in a dollar-for-dollar increase in net foreign liabilities, to approximately $3 trillion. During that year, asset price changes in local currency terms were substantial, but they roughly canceled out, having a minimal impact on the net foreign liabilities of the US. By contrast, the exchange rate valuation effects were substantial. Dollar depreciation that year increased the value of US assets abroad by $416 billion. By year end

2003, the net foreign liabilities of the US were valued at $2.4 trillion dollars, an increase of only $83 billion compared with the previously discussed US capital inflow of $560 billion.

Of course, a real dollar depreciation has a one-off impact on the value of US net foreign assets, and a stabilization of net foreign liabilities as a ratio of US GDP will require a reduction in the ratio of the current account to GDP. However, the current account deficit to GDP ratio need not return to zero for sustainability to be achieved. Indeed, a US current account deficit to GDP ratio in the range of 2 to 3 percent is probably consistent with sustainability at something like the global level of interest rates and equity valuations. Consider this fact: in 2001, US net foreign liabilities were 22.8 percent of US nominal GDP. Two years later, US net foreign liabilities to GDP had risen by a very modest 1.3 percentage points, to 24.1 percent of GDP, notwithstanding current account deficits of roughly 5 percent of GDP in each of 2002 and 2003. The data in Table 3.5 show that exchange rate valuation effects have been important in previous years. For example, in 2002, the exchange rate revaluation of US foreign assets offset 46 percent of the foreign capital inflow; in 1994 and 1995, the exchange rate valuation effect offset 52% of the net capital inflow. Of course, exchange rate appreciation has the opposite effect. Of the $1.3 trillion rise in US net foreign liabilities that accumulated in the three years 1999-2001, $549 billion , or 43 percent, was due to the valuation impact of the appreciation of the dollar that occurred during those years.

Another factor that should be considered when thinking about sustainability and adjustment of international imbalances is the longstanding evidence for the US of substantial differences in the rates of return that US investors earn on their foreign investments com-

pared with the rate of return that foreign investors earn and require on their investments in the US. That is, even though the US is, and has been for many years, the world's largest 'net debtor', with net foreign liabilities estimated to be some $2.4 trillion dollars at year end 2003, the US still to this day earns more interest and dividends on its foreign assets than it pays out on its foreign liabilities, even though the latter exceed the former by more than 2 trillion dollars. Specifically, for 2004, income receipts on US assets abroad totaled $366 billion while income payments on foreign assets in the US totaled $344 billion. How can the US continue to run a surplus on international investment income with its large stock of international liabilities? Differences in portfolio composition can probably account for some of this. For example, in recent years 60 percent of US assets abroad were invested in foreign equities and foreign direct investment. By contrast, only 40 percent of foreign claims against the US were invested in US equities and direct investment. However, in order to account for the persistent surplus in the US international investment income account, portfolio composition is probably not sufficient. In addition, it is likely the case that the US earns consistent higher returns on its FDI than the rest of the world earns on its US FDI.

	1989	1995	1997	1998	1999	2000	2001	2002	2003
Portfolio Shares									
Private US Investment Abroad									
FDI	39.7%	36.8%	36.4%	38.4%	39.6%	37.5%	34.6%	32.0%	35.9%
Securities and Currency	15.0%	32.5%	34.0%	34.6%	35.2%	33.2%	31.6%	29.0%	32.6%
Other Private Assets	45.3%	30.7%	29.6%	27.1%	25.2%	29.3%	33.7%	39.0%	31.5%
Private Foreign Investment in the US									
FDI	26.0%	28.0%	30.7%	34.3%	37.4%	35.0%	31.5%	25.5%	26.9%
Securities and Currency	34.9%	40.9%	42.5%	42.1%	40.6%	41.0%	42.6%	44.6%	47.0%
Other Private Assets	39.1%	31.1%	26.8%	23.6%	22.0%	24.0%	25.9%	30.0%	26.0%

Table 3.6: Portfolio Shares

We see that in both 2003 and 2004, the US earned high returns on FDI, earning profits of 8.7 percent of FDI assets at market value in 2004 and 9.2 percent of FDI assets at market value in 2003. By contrast, foreign owned direct investment assets in the US earned 4.3 percent of assets at market value in 2004 and 3.4 percent of assets at market value in 2003. This disparity is not a recent phenomenon. As the Table shows, the US has consistently since 1989 – the year the US net foreign asset position turned negative – earned higher returns on its FDI assets than foreigners have earned on their US investments. The Table also reports the rate of return on non-FDI assets and liabilities. The absolute return differentials are much smaller, and are consistently negative, indicating that foreign non-FDI holdings pay slightly higher returns than US non-FDI holdings. Once we take into account the differences in portfolio composition between US assets abroad and foreign assets in the US (reported in Table 3.7), we obtain the time series on the total return differential reported in Table 3.6.

	1989	1995	1997	1998	1999	2000	2001	2002	2003	2004
US owned assets abroad										
Total Assets	2,350,235	3,964,558	5,379,128	6,174,518	7,390,427	7,393,643	6,898,707	6,613,320	7,863,968	
US Private Assets	2,094,878	3,703,433	5,158,094	5,941,744	7,169,782	7,180,075	6,683,092	6,369,409	7,595,619	
FDI Assets	832,460	1,363,792	1,879,285	2,279,601	2,839,639	2,694,014	2,314,934	2,039,780	2,730,289	
Foreign Securities	314,294	1,203,925	1,751,183	2,052,995	2,525,341	2,385,353	2,114,734	1,846,879	2,474,374	
Other US Private Assets	948,124	1,135,716	1,527,626	1,609,148	1,804,802	2,100,708	2,253,424	2,482,750	2,390,956	
Income Receipts										
Total Receipts	160270	208065	254534	258871	290474	347614	283761	263861	291354	365886
FDI Receipts	61981	95260	115323	103963	131626	151839	128665	147291	187522	237564
Returns on US owned assets abroad										
Return on all Assets	8.0%	6.3%	5.5%	4.8%	4.7%	4.7%	3.8%	3.8%	4.4%	4.7%
Return on FDI	9.0%	8.5%	7.2%	5.5%	5.8%	5.3%	4.8%	6.4%	9.2%	8.7%
Return on non-FDI assets	7.5%	5.1%	4.6%	4.4%	4.1%	4.3%	3.3%	2.5%	2.3%	2.5%
Foreign-owned assets in the US										
Total Liabilities	2,397,222	4,270,394	6,201,860	7,249,895	8,437,115	8,982,199	9,206,868	9,166,727	10,514,958	
Liabilities to Private Foreigners	2,055,476	3,587,521	5,328,144	6,353,721	7,486,027	7,951,491	8,124,572	7,954,004	9,040,797	
FDI Liabilities	534,734	1,005,726	1,637,408	2,179,035	2,798,193	2,783,235	2,560,294	2,025,345	2,435,539	
Securities and Currency (Cash, US	716,523	1,466,328	2,262,490	2,675,016	3,042,633	3,260,616	3,459,610	3,545,585	4,251,500	
Other Liabilities to Private Foreigne	804,219	1,115,467	1,428,246	1,499,670	1,645,201	1,907,640	2,104,668	2,383,074	2,353,758	
Income Receipts										
Total Payments	-141463	-189353	-244195	-257554	-280037	-329864	-263120	-259626	-261106	-344925
FDI Payments	-7045	-30318	-42950	-38418	-53437	-56910	-12783	-46460	-68657	-105252
Returns on US owned assets abroad										
Return on all Liabilities	7.1%	5.5%	4.9%	4.2%	3.9%	3.9%	2.9%	2.8%	2.8%	3.3%
Return on FDI	1.8%	4.0%	3.5%	2.3%	2.5%	2.0%	0.5%	1.8%	3.4%	4.3%
Return on non-FDI Liabilities	8.4%	5.9%	5.3%	4.8%	4.5%	4.8%	4.0%	3.2%	2.7%	3.0%
Return Differentials										
Total	0.9%	0.8%	0.6%	0.7%	0.8%	0.8%	0.9%	1.0%	1.6%	1.4%
FDI	7.2%	4.5%	3.7%	3.2%	3.3%	3.3%	4.3%	4.5%	5.8%	4.4%
Non-FDI	-0.9%	-0.8%	-0.7%	-0.4%	-0.4%	-0.5%	-0.7%	-0.7%	-0.4%	-0.5%

Table 3.7: US-owned Assets Abroad and Foreign-owned Assets in the US

Another factor that may have delayed adjustment in the US current account is the more modest decline in the broad, real trade weighted dollar as compared with the decline in the dollar that occurred during the 1985-1988. The Federal Reserve's real broad trade weighted dollar index is plotted in Figure 3.7.

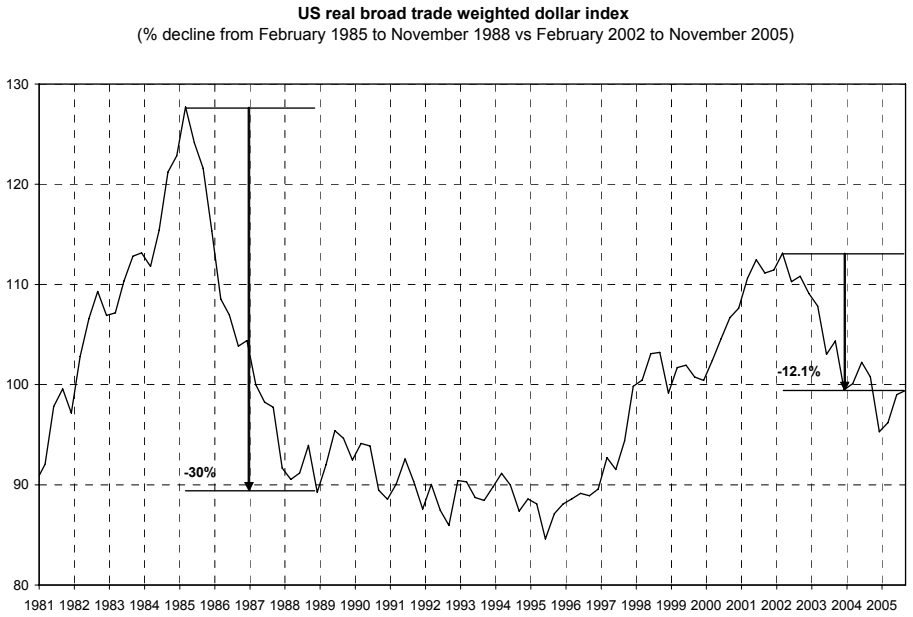


Figure 3.7: Trends in the US Real Effective Exchange Rate

In the three years after the dollar's peak in early 1985, the broad dollar index declined by 30 percent. By contrast, in the three years since the dollar's recent peak in early 2002, it has declined by less than 15 percent. Obviously, the intervention by Asian central banks has limited the depreciation of the dollar against a number of significant US trading partners.

Our final point is that the US current account deficit is in part an endogenous, general equilibrium outcome of global financial and macroeconomic integration. As such, we believe it reflects a global excess supply of saving relative profitable investment opportunities. In a world in which there is a global excess supply of saving relative to investment, we would expect to find and indeed find today that global real interest rates are low and that some country or group of countries must absorb the surplus of internationally mobile

capital. Required real rates return - as measured by yields on TIPS[25] in the US and indexed gilts in the UK - are unusually low (below 2 percent as of this writing). In the late 90s, the opposite was the case and rapid (in retrospect unsustainable) world investment rates surged ahead of savings, pushing up real interest rates (Tips yields were at 4 percent in March 2000 when the bubble peaked). Although no one can say for sure how long the present imbalance between global saving and investment will persist, it seems clear that this global imbalance between saving and investment is contributing to the size of the US current account deficit and its failure to adjust as May 2005.

3.6 Concluding remarks

Are there thresholds of current account adjustment? This chapter has reported evidence in favour of this proposition. We found statistically significant evidence of differing adjustment dynamics in the current-account-to-net-output ratio for all of the G7 countries examined. In particular, each country displayed three regimes — a surplus regime and a deficit regime in which the current account tended to revert towards its long-run mean, albeit at different speeds in each regime (showing that sign does indeed matter), and an 'inertia regime' in which, for intermediate levels of the current account balance between the surplus and deficit regimes, current account adjustment was negligible (showing that size also matters). We also showed, however, that one size does not fit all in the sense that we found significant cross-country variation in the size of the estimated thresholds. We

[25] Treasury Inflation-Protected Securities

also found substantial cross-country variation in the estimated speed of adjustment once countries cross their current account deficit or surplus thresholds.

Our results support the findings of Caroline Freund and Frank Warnock, by providing econometric evidence on the nonlinearities and differences in current account adjustment across industrial countries. In line with their results, countries with large deficits such as the US exhibit relatively wide thresholds within which current account adjustment is absent and relatively slow speeds of adjustment once these thresholds, especially the deficit threshold, are crossed. While our analysis focuses on the relatively homogeneous post Bretton-Woods period, Barry Eichengreen and Muge Adalet present an historical analysis of current account reversals starting from the gold standard period and find evidence of substantial differences in current account adjustments episodes also across time.

We also found evidence of statistically significant shifts in the mean and variance of the probability distribution of several G7 exchange rates, equity prices, and interest rate differentials that occur in conjunction with our estimated current account adjustment regimes. In particular, we found a tendency towards exchange rate depreciation during current account deficit regimes and exchange rate appreciation during current account surplus regimes, and statistically significant increases in exchange rate volatility during current account deficit adjustment regimes for the US, Japan, and Germany. This suggests that a multivariate approach involving the joint modeling of exchange rates and the current account within a nonlinear framework would be a fruitful exercise, as well as being consistent with substantial evidence in favor of nonlinear adjustment in real exchange rates (see, e.g.,

Obstfeld and Taylor (1997); Taylor and Taylor, 2004). This is an avenue we pursue in chapter 5.

Chapter 4
The Savings and Investment Components of Current Account Imbalances

4.1 Introduction

The analysis of the national income identity shows that a current account deficit can be either determined by relatively low savings or high investment levels. The economic literature provides a rationale for the existence of current account deficits in emerging and developing economies given the need of investment for development and growth in these countries. Nevertheless, current global imbalances take place in the form of high deficits in developed countries, notably the US, and surpluses in most emerging markets, such as Asian countries, with substantial capital flows from emerging markets allowing to finance the widening savings-investment gaps.

This chapter aims at disentangling the domestic components of current account imbalances for a group of industrial countries, by highlighting the role of savings and investment on current account adjustment. The first part of the analysis looks at the main stylized facts on private and public savings in G7 countries over the last two decades. We investigate the degree of international capital mobility in these countries and the robustness of the assumption of Ricardian equivalence. The second part of the chapter builds on the methodology and results of chapter 3. In particular we apply the threshold autoregressive methodology to a disaggregated analysis of the components of current account adjustment.

4.2 Stylised facts on private and public savings in the G7

One possible approach for the analysis of current account imbalances is by looking at the national income identity[26]. The study of the national income and product accounts allows to disentangle the different causes of current account surplus/deficit, by highlighting the relationship between public spending and the current account as well as between private savings and investment. In particular, according to the national income identity,

$$Y_t = C_t + I_t + G_t + X_t - M_t + rB_{t-1} \tag{4.14}$$

where Y represents domestic income (GNP), C private consumption, I domestic investment, G government expenditure, $X - M$ net exports and rB is net primary income from abroad, in terms of foreign debt. It follows that the current account can be expressed as

$$Y_t - (C_t + I_t + G_t) = X_t - M_t + rB_{t-1} = CA_t \tag{4.15}$$

If we now look at the national savings identities, we can identify private savings (S^P) as the difference between households' disposable income and consumption, and public savings (S^G) as the difference between government income and government spending. From the national income identity, we can then derive a new definition for total savings and the

[26] See Obstfeld and Rogoff (1994) for a comprehensive textbook analysis of the topic.

current account as follows:

$$
\begin{aligned}
S_t &= S_t^P + S_t^G && (4.16)\\
&= (Y_t - T_t - C_t) + (T_t - G_t) && (4.17)\\
&= Y_t - C_t - G_t && (4.18)\\
&\Rightarrow I_t + CA_t && (4.19)
\end{aligned}
$$

Therefore, in an open economy, savings can be directed either to domestic capital stock or to foreign wealth. Under this definition, the current account balance represents the net foreign investment of the country. In real terms, this implies that excess domestic savings will be matched by positive net exports to foreigners.

Table 4.1 provides a summary of these variables for each of the G7 countries and pinpoints their behaviour across time. Data are presented as percentages of net output, for comparison reasons. We identify surplus countries, such as Japan and Germany, as well as deficit ones, such as the UK and the US. Most countries with a deficit ca position at present show a decrease in private savings in the 2000-2003 period, with some sign of recovery in the latest observations for 2004. In the case of the US, this trend has been partly explained by the effect of high productivity prospects for the US *vis-à-vis* other countries (along seminal work by Glick and Rogoff, 1995), by the steady decline of real interest rates and by the wealth effect arising from the global equity market boom first and the real estate one later (Kraay and Ventura, 2005). In terms of public savings, current account deficits have been associated to government deficits for Italy, the UK and the US.

Savings, Investment and Current Account in G7 Countries

	Year (avgs)	**Public Savings**	**Private Savings**	**Investment**	**Current Account**
		(% of Net Output)			
Canada	1980-1989	-3.62	37.04	35.54	-2.12
	1990-1999	-3.02	30.12	30.67	-3.57
	2000-2003	5.19	30.64	31.95	3.88
	2004	5.12	30.04	33.04	2.12
France	1980-1989	2.72	30.83	36.11	-2.56
	1990-1999	0.59	33.98	32.95	1.62
	2000-2003	7.67	28.90	33.32	3.25
	2004	10.42	24.00	33.88	0.54
Germany	1980-1989	2.10	38.17	34.15	6.13
	1990-1999	1.24	40.42	37.39	4.26
	2000-2003	-0.50	34.63	30.73	3.39
	2004	-2.04	36.10	26.88	7.19
Italy	1980-1989	-10.70	47.04	38.31	-1.97
	1990-1999	-6.66	39.35	31.11	1.58
	2000-2003	0.65	31.66	31.90	0.42
	2004	-0.55	31.57	31.62	-0.60
Japan	1980-1989	10.08	45.47	51.81	3.74
	1990-1999	10.40	45.18	51.20	4.39
	2000-2003	4.12	44.01	43.02	5.12
	2004	3.05	43.98	40.49	6.53
United Kingdom	1980-1989	7.20	21.95	30.73	-1.57
	1990-1999	-1.43	26.13	27.00	-2.30
	2000-2003	1.16	22.56	25.89	-2.17
	2004	-1.76	27.00	25.69	-0.45
United States	1980-1989	-1.28	31.63	31.99	-1.65
	1990-1999	0.22	26.03	27.55	-1.30
	2000-2003	0.93	22.36	29.02	-5.73
	2004	-2.57	24.61	30.33	-8.30

Table 4.1: Savings, Investment and Current Account in G7 Countries

Before moving to a more formal analysis of current account adjustment, we want to look at some of the assumptions behind the national income identity and the resulting current account definitions. In particular, we want to investigate the degree of international capital mobility in these countries and the robustness of the assumption of Ricardian equivalence.

The hypothesis of perfectly integrated capital markets would imply the absence of correlation between domestic savings and investment and as a result between fiscal deficits and investment levels. The seminal 1980 paper by Feldstein and Horioka and more recent work in the literature[27] highlight the weakness of this assumption. Evidence from OECD and developing countries show a significant tendency for savings to be domestically invested. Although the savings retention coefficient has decreased with time and changes in the nature of capital markets, Feldstein(2005) indicates only a modest decline, once countries' GDP weight are accounted for. Among plausible explanations for this apparent puzzle in presence of capital mobility, Obstfeld-Rogoff(1996) suggest the existence of ca targeting strategies by national governments, limited access to capital markets by corporation and therefore a strong impact of corporate savings on investment, the existence of capital endowments close to steady state level of most advanced countries, endogeneity issues related to demographic and productivity changes.

With reference to our country sample, we obtain partial equilibrium results in line with the literature. In particular, we estimate the following simple OLS regressions:

$$\frac{I_t}{Y_t} = \alpha + \beta_s \frac{S_t}{Y_t} \tag{4.20}$$

$$\frac{I_t}{Y_t} = \alpha + \beta_s^G \frac{S_t^G}{Y_t} + \beta_s^P \frac{S_t^P}{Y_t} \tag{4.21}$$

where β_s represents the savings retention coefficient. Results in the first column of Table 4.2 are based on the simple Feldstein-Horioka regression (4.20), while column two and three report results for disaggregated values of savings according to equation (4.21). Esti-

[27] See among others Mussa-Goldstein(1993), Obstfeld-Rogoff (2001), Stulz (2005) and Feldstein (2005).

mates for the savings retention coefficient are around 0.50 for most G7 economies with the exception of Italy and Japan (where the beta coefficients are found not significantly different from 1 at 5% significance level). Although this value is lower than the 0.89 estimate in the original Feldstein-Horioka work, it is still highly significant. Disaggregate savings data, while showing different relationships between public/private savings and investment, support these results. Furthermore, parameter estimates for β_s^G and β_s^P tend to be close in value and, in particular, are not statistically different from each other for France and the US. According to similar results by Feldstein and Bacchetta (1991), these estimates suggest a crowding out effect of government deficits on investment.

Feldstein-Horioka Savings Retention Coefficients

	β_S	βs^G	βs^P
Canada	0.50**	0.40**	0.70**
	(0.060)	(0.059)	(0.070)
France	0.51**	0.55**	0.46**
	(0.111)	(0.119)	(0.121)
Germany	0.59**	1.07**	0.43**
	(0.083)	(0.164)	(0.092)
Italy	1.14**	0.70**	0.96**
	(0.090)	(0.117)	(0.087)
Japan	1.09**	1.09**	1.35**
	(0.050)	(0.048)	(0.094)
United Kingdom	0.55**	0.54**	0.36**
	(0.076)	(0.074)	(0.106)
United States	0.56**	0.60**	0.54**
	(0.046)	(0.042)	(0.056)

Note: SE in parenthesis.**, * indicate significance at 1% and 5% level, respectively.

Table 4.2: Feldstein-Horioka Savings Retention Coefficients

In addition, we find that cross-correlations between the private-savings series and the public-savings one have a negative sign but a value lower than one, as reported in Table 4.3. Under full Ricardian equivalence, a government should have no effect on overall savings of the economy and just a temporary one in special cases, such as a debt financed increase in spending. An increase of private savings would in fact match the reduced level of public savings. The current account balance would be thus left unchanged. Our estimates suggest a much weaker link between public and private savings and therefore imply a relaxation of the Ricardian Equivalence and, as a result, account for the possibility of a twin deficit (both public and external) fed by a deterioration of the fiscal stance. In particular, these conclusions are consistent with the imperfect capital mobility results identified in the previous analysis.

Cross-Correlations of Private and Public Savings

Lags:	0	1	2	3	4
Canada	-0.501	-0.413	-0.338	-0.285	-0.236
France	-0.707	-0.650	-0.613	-0.566	-0.532
Germany	0.035	0.087	0.138	0.196	0.260
Italy	-0.918	-0.882	-0.848	-0.810	-0.778
Japan	-0.481	-0.382	-0.288	-0.211	-0.148
United Kingdom	-0.737	-0.733	-0.731	-0.698	-0.637
United States	-0.431	-0.442	-0.467	-0.471	-0.472

Table 4.3: Cross-correlations of Private and Public Savings

4.3 Thresholds of current account adjustment

In chapter 3, we report evidence in favour of the existence of nonlinearities of current account adjustment. We find statistically significant evidence of differing adjustment dynamics in the current-account-to-net-output ratio for all of the G7 countries. In particular, we model the nonlinearities in the ca by means of an equilibrium threshold autoregressive (TAR) model. The baseline specification used in the analysis is as follows:

$$\begin{aligned} \mathbf{ca}_t &= \overline{\rho}\times\mathbf{1}\{\mathbf{ca}_{t-d},\overline{\delta}\}\times\mathbf{ca}_{t-1}+\underline{\rho}\times\mathbf{1}\{\mathbf{ca}_{t-d},\underline{\delta}\}\times\mathbf{ca}_{t-1}+ \\ &\quad +(\mathbf{1}-\mathbf{1}\{\mathbf{ca}_{t-d},\overline{\delta}\})\times(\mathbf{1}-\mathbf{1}\{\mathbf{ca}_{t-d},\underline{\delta}\})\times\mathbf{ca}_{t-1}+\mathbf{e}_t \end{aligned} \tag{4.22}$$

where $1\{ca_{t-d},\overline{\delta}\}$ is an indicator function that takes on a value of 1 when $ca_{t-d} > \overline{\delta} \geqslant 0$ (and zero otherwise) and $1\{ca_{t-d},\underline{\delta}\}$ is an indicator function that takes on a value of 1 when $ca_{t-d} < \underline{\delta} \leqslant 0$ (and zero otherwise). This approach postulates that the persistence of the current account imbalance in a country may depend upon whether or not the current account imbalance has crossed a surplus 'threshold' of $\overline{\delta} \geqslant 0$ or a deficit threshold of $\underline{\delta} \leqslant 0$.

We show how each G7 country does in fact display three regimes — a surplus regime and a deficit regime in which the current account tend to revert towards its long-run mean, albeit at different speeds in each regime, and an 'inertia regime' in which, for intermediate levels of the current account balance between the surplus and deficit regimes, current account adjustment is negligible. We can identify significant cross-country variation in the size of the estimated thresholds as well as substantial cross-country variation in the es-

timated speed of adjustment once countries cross their current account deficit or surplus thresholds.

4.3.1 Disentangling the domestic components of current account imbalances

In this section we plan to build on the analysis of chapter 3, summarised above, and decompose the speed of ca adjustment of the univariate model in its saving and investment components, in order to identify the effect of changes in these variables on current account imbalances. Specifically, we look at disaggregated data for savings, by analysing ca adjustment in terms of both Government and Private sector savings. We adopt the following specification:

$$\begin{aligned} ca_t &= \overline{\alpha} \times 1\{ca_{t-d}, \overline{\delta}\} \times s^G_{t-1} + \underline{\alpha} \times 1\{ca_{t-d}, \underline{\delta}\} \times s^G_{t-1} + \\ &+ \overline{\beta} \times \mathbf{1}\{ca_{t-d}, \overline{\delta}\} \times \mathbf{s}^P_{t-1} + + \underline{\beta} \times \mathbf{1}\{ca_{t-d}, \underline{\delta}\} \times \mathbf{s}^P_{t-1} + \\ &+ \overline{\gamma} \times \mathbf{1}\{ca_{t-d}, \overline{\delta}\} \times \mathbf{i}_{t-1} + \underline{\gamma} \times \mathbf{1}\{ca_{t-d}, \underline{\delta}\} \times \mathbf{i}_{t-1} + \\ &+ (1 - 1\{ca_{t-d}, \overline{\delta}\}) \times (1 - 1\{ca_{t-d}, \underline{\delta}\}) \times ca_{t-1} + \mathbf{v}_t \end{aligned} \tag{4.23}$$

We look at quarterly data for the period 1979:1 to 2004:4. The data in the analysis are obtained from the International Financial Statistics (IFS) Database by the IMF. All variables are seasonally adjusted and expressed in national currency. According to the national account statistics, the current account variable is estimated as the sum of net exports and net primary income from abroad (see above, [4.15]); net output is obtained by subtract-

ing Government consumption expenditure and gross fixed capital formation (investment) to GDP.

In line with the univariate results of chapter 3, the two asymmetric thresholds in the TAR model are selected jointly by minimisation of the overall sum of squared errors. The estimation method involves a double grid search over ca. Following Hansen (1997), the range for the grid search is selected a priori to contain ca observations in between the 15th ($\underline{ca}$) and the 85th percentile ($\overline{ca}$). This reduction in the grid range is needed in order to avoid sorting too few observations in one regime for extreme values of the thresholds. As a result, the appropriate ranges are defined as $\overline{\mathbf{R}} = [\mu, \overline{ca}]$ and $\underline{\mathbf{R}} = [\mu, \underline{ca}]$, for $\overline{\delta}$ and $\underline{\delta}$ respectively. We select a delay equal to 2.

As the minimisation process for a three-regime/ two-threshold TAR process is numerically intensive, we rely on the estimation methodology proposed by Hansen (1999) for multiple thresholds. This consists of a three-stage grid search, where the second-stage estimation of the two-threshold model is made conditional on the first-stage single-threshold estimate of δ (either $\overline{\delta}$ or $\underline{\delta}$), the third stage being used as a refinement.

Threshold models of de-meaned CA/NO wrt Government S/NO, Private S/NO and I/NO

Q1 1979- Q4 2004	Thresholds of CA/NO		Model estimation		
COUNTRY	Upper threshold	Lower threshold		ca	
			1(ca{2}>T1)*s^{G}{1}	0.660	(0.090)
	2.95	**-4.32**	1(ca{2}<T2)*s^{G}{1}	0.713	(0.179)
			1(ca{2}>T1)*s^{P}{1}	0.937	(0.061)
CANADA			1(ca{2}<T2)*s^{P}{1}	0.944	(0.064)
			1(ca{2}>T1)*i{1}	-0.589	(0.146)
			1(ca{2}<T2)*i{1}	-0.924	(0.148)
			1(T1>ca{2}>T2)*ca{1}	1.000	
			1(ca{2}>T1)*s^{G}{1}	0.782	(0.097)
	3.67	**-3.11**	1(ca{2}<T2)*s^{G}{1}	0.771	(0.122)
			1(ca{2}>T1)*s^{P}{1}	0.933	(0.130)
FRANCE			1(ca{2}<T2)*s^{P}{1}	0.618	(0.203)
			1(ca{2}>T1)*i{1}	-1.097	(0.085)
			1(ca{2}<T2)*i{1}	-1.035	(0.073)
			1(T1>ca{2}>T2)*ca{1}	1.000	
			1(ca{2}>T1)*s^{G}{1}	0.755	(0.084)
	0.42	**-0.45**	1(ca{2}<T2)*s^{G}{1}	0.784	(0.081)
			1(ca{2}>T1)*s^{P}{1}	0.958	(0.172)
GERMANY			1(ca{2}<T2)*s^{P}{1}	0.452	(0.213)
			1(ca{2}>T1)*i{1}	-0.944	(0.084)
			1(ca{2}<T2)*i{1}	-0.936	(0.088)
			1(T1>ca{2}>T2)*ca{1}	1.000	
			1(ca{2}>T1)*s^{G}{1}	0.702	(0.102)
	0.00	**-0.19**	1(ca{2}<T2)*s^{G}{1}	0.529	(0.118)
			1(ca{2}>T1)*s^{P}{1}	0.655	(0.110)
ITALY			1(ca{2}<T2)*s^{P}{1}	0.550	(0.112)
			1(ca{2}>T1)*i{1}	-0.916	(0.071)
			1(ca{2}<T2)*i{1}	-0.728	(0.089)
			1(T1>ca{2}>T2)*ca{1}	1.000	
			1(ca{2}>T1)*s^{G}{1}	1.071	(0.178)
	1.63	**-1.08**	1(ca{2}<T2)*s^{G}{1}	0.776	(0.150)
			1(ca{2}>T1)*s^{P}{1}	0.915	(0.082)
JAPAN			1(ca{2}<T2)*s^{P}{1}	0.957	(0.073)
			1(ca{2}>T1)*i{1}	-0.979	(0.068)
			1(ca{2}<T2)*i{1}	-0.855	(0.061)
			1(T1>ca{2}>T2)*ca{1}	1.000	
			1(ca{2}>T1)*s^{G}{1}	0.264	(0.115)
	0.03	**-0.29**	1(ca{2}<T2)*s^{G}{1}	0.601	(0.164)
			1(ca{2}>T1)*s^{P}{1}	0.506	(0.104)
UK			1(ca{2}<T2)*s^{P}{1}	0.712	(0.136)
			1(ca{2}>T1)*i{1}	-0.725	(0.192)
			1(ca{2}<T2)*i{1}	-0.851	(0.097)
			1(T1>ca{2}>T2)*ca{1}	1.000	
			1(ca{2}>T1)*s^{G}{1}	0.852	(0.078)
	0.46	**-1.58**	1(ca{2}<T2)*s^{G}{1}	1.056	(0.172)
			1(ca{2}>T1)*s^{P}{1}	0.787	(0.067)
US			1(ca{2}<T2)*s^{P}{1}	0.880	(0.196)
			1(ca{2}>T1)*i{1}	-0.755	(0.090)
			1(ca{2}<T2)*i{1}	-0.605	(0.242)
			1(T1>ca{2}>T2)*ca{1}	1.000	
				SE in brackets	

Specification:

*ca = a1*1(ca{2}>T1)*sG{1} + a2*1(ca{2}<T2)*sG{1} + b1*1(ca{2}>T1)*sP{1} + b2*1(ca{2}<T2)*sP{1} + c1*1(ca{2}>T1)*i{1} + c2*1(ca{2}<T2)*i{1} + 1(T1>ca{2}>T2)*(sG+sP-i){1} + e*

Table 4.4: Threshold model of CA/NO and its domestic components

Results are reported in Table 4.4. Variables enter with the correct sign suggested by the national accounting relationship, that is with a negative effect of lagged investment on ca and positive of private and public savings. We find cross-country variation in the

threshold values, which range for surplus regimes from the 3.67% of ca of France to the immediate correction of Italy and for deficit regimes from the -4.32% level of Canada to again a low threshold value of -0.19% for Italy. As regards the effect of the savings and investment components on ca, deficit persistent countries, that is Canada, the UK and the US, present very significantly higher slope parameters in the deficit regime, especially in the case of lagged savings. The opposite can be said for surplus countries.

In terms of public and private savings, for the UK and Canada the private component seems to have affected the slow ca adjustment from its deficit regime. On the contrary, in the US, government savings have a predominant role in the sluggishness of the ca variable, supporting the idea of a twin deficits where the poor fiscal stance feeds into the current account.

4.4 Concluding remarks

In this chapter, we look at the national income identity as a way to decomposing the current account into its domestic components. We analyse trends in savings (both public and private) and investment in G7 countries over the last two decades and find evidence that most deficit countries have experienced declining private savings in the last years and negative levels of public savings, especially in the case of Italy, the UK and the US.

A simple analysis of the main assumptions behind the national income identity and the resulting current account definitions shows significant savings retention coefficients, that is the existence, despite integrated capital markets, of a correlation between domestic savings and investment, with a crowding out effect of government deficits on investment.

Consistently with these results, we also find evidence of a correlation lower than minus one between public and private savings, against the Ricardian hypothesis that government spending should have no long run effect on the overall savings of the economy and thus on the current account.

Given these preliminary insights, we investigate the role of savings and investment in current account adjustment by extending the threshold methodology of chapter 3 to a more disaggregated framework. Our estimates present the expected signs, as suggested by the national accounting relationship, with a negative effect of lagged investment on ca and positive of private and public savings. In line with our previous results, we find cross-country variation in the threshold values. Deficit countries present significantly persistent slope coefficient in the deficit regime, especially in the case of lagged savings. For the UK and Canada, sluggishness in ca adjustment from its deficit regime seems mainly due to the private savings component. For the US, the slow speed of adjustment of the government savings component supports the hypothesis of a twin deficits (both public and external) fed by a deterioration of the fiscal stance.

While the analysis of this chapter tries to shed some light on the domestic factors affecting ca adjustment in G7 countries, its conclusions are not exhaustive. A comprehensive explanation of the current global imbalances would require an analysis of global developments in savings patterns across the world, with a special emphasis to the increase in global savings led by emerging markets (what Chairman Bernanke defines as the "savings glut") and to productivity levels, rates of return and asset prices in the US and other deficit countries.

Chapter 5
Thresholds of Real Exchange Rate and Current Account Adjustment

5.1 Introduction

While the theoretical literature has been widely focusing on the role of relative prices, that is of real exchange rates, in current account adjustment[28], there are only few empirical works trying to produce estimates of this relationship. The recent empirical contribution by Chinn and Lee (2005) has tried to fill this gap by looking at a VAR model of the exchange rate and current account. Furthermore, the strict link between exchange rate and *ca* adjustment has been further corroborated by recent empirical research on the role of the valuation channel in current account adjustments, in particular in the part accounted for by changes in the exchange rate (see Gourinchas and Rey (2005), Lane and Milesi-Ferretti (2005), Freund and Warnock (2005)).

Time series analysis of the real exchange rate and current account has identified in the last years evidence of nonlinear behaviour in both these variables. To date, however, there is no comprehensive explanation of their joint dynamics, in particular on the role of the real exchange rate in determining the nonlinear adjustment of the current account.

Since the seminal work by Obstfeld and Taylor (1997) and Taylor et al. (2001) the literature on the real exchange rate has provided evidence of strong nonlinearities in its

[28] See for example the latest works by Obstfeld and Rogoff (2004).

dynamics, showing that real exchange rates are stationary but exhibit nonlinear adjustment to their unconditional mean. In chapter 3, we reported statistically significant evidence in favour of differing adjustment dynamics in the current-account-to-net-output ratio for the G7 countries. These results supported the point, shared by several observers[29], that in the current global economic environment there is a "threshold" current account imbalance beyond which adjustment must ultimately take place, even if evidence of adjustment is scarce or non-existent before the threshold is reached.

Furthermore, we partially addressed the interlinkages between real exchange rate and current account adjustment, by means of a simple GARCH analysis and found evidence of statistically significant shifts in the mean and variance of the probability distribution of several G7 exchange rates, equity prices, and interest rate differentials that occurred in conjunction with the estimated current account adjustment regimes. In particular, we found a tendency towards exchange rate depreciation during current account deficit regimes and exchange rate appreciation during current account surplus regimes, and statistically significant increases in exchange rate volatility during current account deficit adjustment regimes for the US, Japan, and Germany. This suggested scope for further research in the joint modeling of exchange rates and the current account within a nonlinear framework, which is the purpose of this chapter. Extending the empirical TAR methodology to a bivariate context for all the G7 countries, we find evidence of a strict link between current account adjustments and deviations of a country real exchange rate from its long run equilibrium, such

[29] See Bergsten (2002), Holman (2001), Greenspan (2003).

that beyond a certain appreciation/depreciation level of the real exchange rate, a country CA imbalance would start reverting towards its mean value.

The plan of the chapter is as follows. Section II estimates and tests a threshold model of the exchange rate in G7 countries, confirming previous results in the literature. Section III conducts a similar exercise for the current account, along the lines of the analysis in chapter 3. Section IV estimates first an univariate CA model with thresholds of reer and then moves to the estimation of a bivariate threshold model for the analysis of the current account and real exchange rate dynamics. Section VI identifies the speeds of current account adjustment in G7 countries, by means of nonlinear impulse response analysis. Section VII concludes.

5.2 Estimating and testing threshold models of the exchange rate

In the literature on the exchange rate and its predictability, there is strong empirical evidence of threshold effects in real exchange rate adjustment (Obstfeld and Taylor, 1997; Taylor, Peel and Sarno, 2001; Kilian and Taylor, 2003; Imbs et al., 2003; Taylor and Taylor, 2004). In particular, these authors show that real exchange rates are stationary but exhibit nonlinear adjustment to their unconditional mean.

Building on these contributions, we estimate and test for each G7 country a threshold autoregressive model of the real effective exchange rate (REER) using the univariate approach developed in Hansen (1996)[30]. We allow for and estimate country-specific

[30] We initially test for non-linearity in G7 current account/net output adjustment following the non-parametric test for nonlinearity developed by Luukkonen et al. (1988) and Terasvirta (1994). Evidence of nonlinear ad-

means, country and regime-specific thresholds, and country and regime specific dynamic adjustment once the exchange has crossed either of the thresholds. Letting $q = 100 * (\log(REER) - \mu)$, we write the equilibrium threshold autoregressive (TAR) model as

$$\begin{aligned} \mathbf{q}_t &= \overline{\rho}\times\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\} \times \mathbf{q}_{t-1}+\underline{\rho}\times\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\} \times \mathbf{q}_{t-1} + \\ &\quad +(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\}) \times (\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\}) \times \mathbf{q}_{t-1}+\mathbf{e}_t \end{aligned} \tag{5.24}$$

where $1\{q_{t-d}, \overline{\delta}\}$ is an indicator function that takes on a value of 1 when $q_{t-d} > \overline{\delta} \geqslant 0$ (and zero otherwise) and $1\{q_{t-d}, \underline{\delta}\}$ is an indicator function that takes on a value of 1 when $q_{t-d} < \underline{\delta} \leqslant 0$ (and zero otherwise). This approach postulates that the persistence of an exchange rate appreciation (depreciation) in a country may depend upon whether or not the exchange rate has crossed an upper threshold of $\overline{\delta} \geqslant 0$ or a lower threshold of $\underline{\delta} \leqslant 0$. We note that a special case of the threshold model is the case in which $\overline{\delta} = \underline{\delta} = 0$ and $\overline{\rho} = \underline{\rho} < 1$ in which case it collapses to a linear stationary AR(1) process. We experimented with a threshold TAR(2) specification but found in general the second lag terms to be insignificant, and thus confine our presentation to the TAR(1) models. We also select a delay parameter d of two quarters as this maximises the fit of the regression in each case.

The threshold model can potentially identify three regimes of real exchange rate adjustment: an upper adjustment regime, a lower adjustment regime, and an 'inertia' regime $\overline{\delta} < \mathbf{q}_{t-2} < \underline{\delta}$ in which the real exchange rate appears to follow a random walk. In a more general smooth threshold transition autoregressive or STAR model (e.g. Taylor, Peel and Sarno, 2001), the speed of adjustment does not increase discontinuously at the threshold;

justment is indicated at the 5% significance level for all countries.

rather, the further way is the exchange rate from its long-run mean, the faster it adjusts. Although STAR models proved useful tools in order to identify the nonlinear behaviour of the exchange rate, we prefer to adopt a simple TAR model, given evidence on the discrete nature of the current account adjustment and our interest in investigating exchange rate and current account adjustment in a multivariate framework.

Before presenting the results, we will discuss some issues involved in the estimation and testing of these model for a system comprised of the G7 countries. The $\mathbf{q}$ variables for the G7 group are first demeaned. A non-zero mean proves to be applicable for all G7 countries. In particular, we detect a structural break in the German series in 1991, corresponding with the German unification; we account for the break by allowing two different means in the real exchange rate for the pre- and post-unification periods.

The two asymmetric thresholds in the TAR model are selected jointly by minimisation of the overall sum of squared errors. The estimation method involves a double grid search over $\mathbf{q}$. Following Hansen (1997), the range for the grid search is selected a priori to contain $\mathbf{q}$ observations in between the 15th ($\underline{\mathbf{q}}$) and the 85th percentile ($\overline{\mathbf{q}}$). This reduction in the grid range is needed in order to avoid sorting too few observations in one regime for extreme values of the thresholds. As a result, the appropriate ranges are defined as $\overline{\mathbf{R}} = [\mu, \overline{\mathbf{q}}]$ and $\underline{\mathbf{R}} = [\mu, \underline{\mathbf{q}}]$, for $\overline{\delta}$ and $\underline{\delta}$ respectively.

As the minimisation process for a three-regime/ two-threshold TAR process is numerically intensive, we rely on the estimation methodology proposed by Hansen (1999) for multiple thresholds. This consists of a three-stage grid search, where the second-stage es-

timation of the two-threshold model is made conditional on the first-stage single-threshold estimate of δ (either $\overline{\delta}$ or $\underline{\delta}$), the third stage being used as a refinement.

Furthermore, final estimates of slope parameters and standard errors for the G7 group of countries are obtained by seemingly unrelated regression (SUR) estimation, in order to allow for potential correlation between the disturbances of the different $\mathbf{q}$ equations, due to common unobservable factors.

Once the thresholds have been selected, according to standard asymptotic theory, (1) is linear in the parameters. As with any simple dummy-variable regression, it can be estimated by linear methods. However, statistical inference in a TAR model bears the difficulty that the thresholds $\overline{\delta}$ and $\underline{\delta}$ may be not identified under the null hypothesis in question (Davies, 1987). In this case, the usual χ^2 distribution needs to be replaced by an approximated empirical distribution obtained by bootstrapping the residuals (Hansen 1997). In particular, artificial observations are calibrated using the restricted estimates and are then used to obtain new estimates of the restricted and unrestricted model (for an application, see Peel and Taylor 2002). The percentage of bootstrap samples – we run 1000 replications – for which the simulated likelihood-ratio statistics exceeds the actual one forms the bootstrap approximation to the p-value of the test statistic under question.

The estimation and testing results are presented in Table 5.1. First the test results: when we test the null hypothesis a single threshold for all countries versus the alternative hypothesis of two thresholds, we reject the null hypothesis in favor of the alternative. This is consistent with three regimes for each country. Second, when we test the hypothesis that the current account follows a random walk inside the 'inertia' regime against the alternative

that it follows a mean reverting autoregressive process inside the inertia regime (a more general formulation of the threshold model) we are unable to reject the null of a random walk inside the inertia regime. In summary, the statistical tests find evidence of non-linear reer adjustment and also identify significant thresholds beyond which this adjustment takes place.

Threshold models of de-meaned log REER

Q1 1979- Q4 2004	Thresholds of REER (asymmetric band)		Slope coefficients (estimation by SUR)			Means
COUNTRY	Upper threshold	Lower threshold	above	band	below	REER
CANADA	**10.62**	**-12.26**	0.939 (0.026)	1.000	0.941 (0.030)	**473.31**
FRANCE	**0.42**	**-4.54**	0.887 (0.037)	1.000	0.911 (0.039)	**467.97**
GERMANY	**1.85**	**-4.30**	0.853 (0.036)	1.000	0.918 (0.036)	**473.76** **468.46**
ITALY	**10.61**	**-6.22**	0.927 (0.037)	1.000	0.877 (0.045)	**466.83**
JAPAN	**12.09**	**-0.36**	0.906 (0.026)	1.000	0.913 (0.018)	**438.15**
UK	**10.29**	**-9.13**	0.874 (0.052)	1.000	0.866 (0.041)	**448.16**
US	**15.26**	**-10.37**	0.960 (0.020)	1.000	0.920 (0.025)	**453.32**

SE in brackets

Bootstrap: *LR-test for band coefficient equal to 1 (SUR): marg.signif. level = 0.994*
LR-test for single threshold (SUR): marg. signif. level = 0.000

Table 5.1: Threshold Models of REER

In the results reported in Table 5.1, we find evidence of threshold behaviour in the real effective exchange rate, with wide thresholds and parameter variations across G7 countries. The inertia band is delimited in between 10.62 and -12.26 for Canada, where is just in between 0.42 and -4.54 for France. As regards the parameter estimates, Canada and US present the slowest speed of adjustment, both above and below the thresholds.

5.3 ”Are there thresholds of current account adjustment?”: an update

In chapter 3, we reported evidence in favour of this proposition. By means of a TAR model similar to that one described in the previous section for the real effective exchange rate, we found statistically significant evidence of differing adjustment dynamics in the current-account-to-net-output ratio for all of the G7 countries examined. In particular, each country displayed three regimes — a surplus regime and a deficit regime in which the current account tended to revert towards its long-run mean, albeit at different speeds in each regime (showing that sign does indeed matter), and an ‘inertia regime’ in which, for intermediate levels of the current account balance between the surplus and deficit regimes, current account adjustment was negligible (showing that size also matters). We also showed, however, that one size does not fit all in the sense that we found significant cross-country variation in the size of the estimated thresholds. We also found substantial cross-country variation in the estimated speed of adjustment once countries cross their current account deficit or surplus thresholds.

Threshold models of de-meaned CA/NO

Q1 1979- Q4 2004	Thresholds (asymmetric band)		Slope coefficients (estimation by SUR)			Means	
COUNTRY	Upper threshold	Lower threshold	above	band	below	'Surplus'	'Deficit'
CANADA	**1.20**	**-4.25**	0.912 (0.048)	1.000	0.927 (0.060)		**-1.587**
FRANCE	**3.64**	**-1.11**	0.964 (0.047)	1.000	0.888 (0.043)	**0.244**	
GERMANY	**0.31**	**-0.31**	0.857 (0.075)	1.000	0.862 (0.064)	**6.185** Pre-1991 **3.536** Post-1991	
ITALY	**3.43**	**0.00**	0.895 (0.083)	1.000	0.771 (0.071)	**0.008**	
JAPAN	**0.80**	**-0.31**	0.924 (0.053)	1.000	0.918 (0.035)	**4.072**	
UK	**0.00**	**-0.29**	0.433 (0.101)	1.000	0.840 (0.097)		**-1.911**
US	**2.37**	**-2.54**	0.927 (0.036)	1.000	**1.051** (0.031)		**-2.237**

SE in brackets

Bootstrap: *LR-test for band coefficient equal to 1 (SUR): marg.signif. level = 0.544*
LR-test for single threshold (SUR): marg. signif. level = 0.003

Table 5.2: Threshold Models of CA/NO: 1979 - 2004

In Table 5.2 we replicate those results extending the sample period to end- 2004[31]. Results are comparable to the previous one, even if we do notice differences in the band width and in the adjustment coefficients. For example, the addition of more recent observation narrows the inertia band for Germany, which present a less persistent surplus regime and an upper threshold closer to its long run surplus mean and symmetric with respect to the lower threshold. Once again, results on the US are a case of their own, with a deficit adjustment coefficient of 1.051 (whereas the sample ending in 2003 estimated it at 0.973) and

[31] Tests for non-linearity - following the Terasvirta(1994) methodology - on the updated current account series show evidence of nonlinear adjustment at 5% significance level for France, Germany, Japan and the US, and at 10% for the UK.

a lower threshold of -2.54. This result reflects the delayed adjustment of .the US current account in recent years, which is at odds with evidence from previous periods.

Nevertheless, while an univariate analysis of current account adjustment in G7 countries can provide useful insights on the differing dynamics and adjustment magnitudes affecting each country, it only represents a first step in the analysis as it fails to provide an explanation behind these specific patterns.

In the previous chapter, we partially addressed this issue, by means of a simple GARCH analysis and found evidence of statistically significant shifts in the mean and variance of the probability distribution of several G7 exchange rates, equity prices, and interest rate differentials that occur in conjunction with our estimated current account adjustment regimes. In particular, we found a tendency towards exchange rate depreciation during current account deficit regimes and exchange rate appreciation during current account surplus regimes, and statistically significant increases in exchange rate volatility during current account deficit adjustment regimes for the US, Japan, and Germany. This suggested scope for further research in the joint modeling of exchange rates and the current account within a nonlinear framework.

5.4 What is determining the CA adjustment? The role of the real exchange rate

In the previous sections, we confirmed evidence in the literature that both the real exchange rate and the current account in G7 countries are stationary but exhibit nonlinear adjustment to their unconditional mean. In this section of the chapter, we plan to explore the empirical

relationship between thresholds of real exchange rate adjustment and thresholds of current account adjustment. In fact, in line with our previous results, we would expect to find a strict link between current account adjustments and deviations of a country real exchange rate from its long run equilibrium, such that beyond a certain appreciation/depreciation of the real exchange rate, a country CA imbalance would start reverting towards its mean value.

Our approach is consistent with the recent theoretical contributions by Obstfeld and Rogoff (2004) and the empirical VAR approach adopted by Chinn and Lee (2005) in emphasizing the jointly determination of exchange rate adjustments and reductions on tradable goods consumption. Furthermore, evidence on the link between exchange rate and *ca* adjustment seems to corroborate recent empirical evidence on the role of the valuation channel in current account adjustments, in particular in the part accounted for by changes in the exchange rate (see Gourinchas and Rey (2005), Lane and Milesi-Ferretti (2005), Freund and Warnock (2005)). However, differently from those valuable cross-sectional investigations, the trade and valuation channels cannot be disentangled in our analysis due to limited data availability on net foreign assets for time series analysis[32].

Letting $ca = 100 * CA/NO$, we estimate a new equilibrium threshold autoregressive model for the ca variable as

[32] Lane and Milesi-Ferretti(2005) provides an extensive database on net foreign assets at annual frequency. The use of linearly interpolated quarterly data does not represent a feasible solution as it would bias the nonlinear nature of the data.

$$\begin{aligned} ca_t &= \overline{\rho}\times\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\}\times\mathbf{ca}_{t-1}+\underline{\rho}\times\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\}\times\mathbf{ca}_{t-1}+ \\ &\quad +(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\})\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\})\times\mathbf{ca}_{t-1}+\mathbf{e}_t \end{aligned} \tag{5.25}$$

This approach postulates that the persistence of a current account surplus (deficit) in a country may depend upon whether or not the real exchange rate has crossed an upper threshold of $\overline{\delta} \geqslant 0$ or a lower threshold of $\underline{\delta} \leqslant 0$, i.e. on the magnitude of its appreciation (depreciation). Because the exchange rate is an asset price, and because of the well documented lags between real exchange rate and current account adjustment, we allow for a delay of 2 quarters in the current account equation.

Estimates for country-specific means, country and regime-specific thresholds, and country and regime specific autoregressive dynamics are reported in Table 5.3[33]. We find strong evidence on the existence of thresholds of reer in ca adjustments in the G7 country sample. We notice once again wide cross-country variation in the estimated adjustment thresholds. For example, the estimated current account deficit adjustment for the US takes place for deviations of the real exchange rate from its long run mean of -11%, and for Japan for over -27%, while for France for just a -0.6% correction. We estimate a similar variation also for surplus adjustments, whereas the UK presents one of the highest reer thresholds among G7 countries with a deviation from equilibrium of 8% before CA adjustment, while Germany presents an immediate adjustment towards its mean. We also see substantial cross-country variation in the estimated autoregressive dynamics once countries cross

[33] Note that CA/NO means differ from those ones in Table 4.2 due to the different data availability for the reer variable (from 1995:1).

their exchange rate thresholds. For deficit adjustment episodes, the estimated autoregressive coefficients range from 0.621 for the UK to 0.941 for the US. For surplus adjustment episodes, the estimated autoregressive coefficients range from 0.945 in the US to 0.417 in the UK. Note, in particular, the case of France where the surplus coefficient does not seem to adjust to a reer appreciation.

Threshold models of de-meaned CA/NO

CA/NO Q1 1979- Q4 2004	Thresholds of REER (asymmetric band)		Slope coefficients (estimation by SUR)			Means	
COUNTRY	Upper threshold	Lower threshold	above	band	below	CA/NO	REER
CANADA	**1.73**	**-13.19**	0.915 (0.051)	1.000	0.832 (0.076)	**-1.519**	**473.31**
FRANCE	**3.70**	**-0.58**	1.148 (0.063)	1.000	0.853 (0.041)	**0.244**	**467.97**
						Pre-/Post 1991	
GERMANY	**0.00**	**-5.65**	0.869 (0.071)	1.000	0.624 (0.139)	**6.576** **3.536**	**473.76** **468.46**
ITALY	**1.68**	**--5.92**	0.700 (0.115)	1.000	0.853 (0.073)	**-0.080**	**466.83**
JAPAN	**4.48**	**-27.06**	0.773 (0.073)	1.000	0.771 (0.044)	**4.289**	**438.15**
UK	**8.04**	**-2.82**	0.417 (0.146)	1.000	0.621 (0.104)	**-1.965**	**448.16**
US	**3.64**	**-10.98**	0.945 (0.032)	1.000	0.941 (0.059)	**-2.350**	**453.32**

SE in brackets

Table 5.3: CA/NO Models with Thresholds of REER

Given the encouraging results of the univariate CA model with thresholds of reer, we proceed with the estimation of a threshold model for the analysis of the current account and real exchange rate dynamics, by treating both variables endogenously as in the following bivariate system:

$$
\begin{aligned}
ca_t \;=\; & \overline{\alpha}\times\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\}\times\mathbf{ca}_{t-1}+\underline{\alpha}\times\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\}\times\mathbf{ca}_{t-1}+\\
& +\overline{\underline{\alpha}}\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\})\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\})\times\mathbf{ca}_{t-1}+\\
& +\overline{\beta}\times\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\}\times\mathbf{q}_{t-1}+\underline{\beta}\times\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\}\times\mathbf{q}_{t-1}+\\
& +\overline{\underline{\boldsymbol{\beta}}}\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\})\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\})\times\mathbf{q}_{t-1}+\mathbf{e}_t
\end{aligned}
\tag{5.26}
$$

$$
\begin{aligned}
q_t \;=\; & \overline{\gamma}\times\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\}\times\mathbf{ca}_{t-1}+\underline{\gamma}\times\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\}\times\mathbf{ca}_{t-1}+\\
& +\overline{\underline{\gamma}}\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\})\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\})\times\mathbf{ca}_{t-1}+\\
& +\overline{\delta}\times\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\}\times\mathbf{q}_{t-1}+\underline{\delta}\times\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\}\times\mathbf{q}_{t-1}+\\
& +\overline{\underline{\boldsymbol{\delta}}}\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\overline{\delta}\})\times(\mathbf{1}-\mathbf{1}\{\mathbf{q}_{t-d},\underline{\delta}\})\times\mathbf{q}_{t-1}+\mathbf{u}_t
\end{aligned}
\tag{5.27}
$$

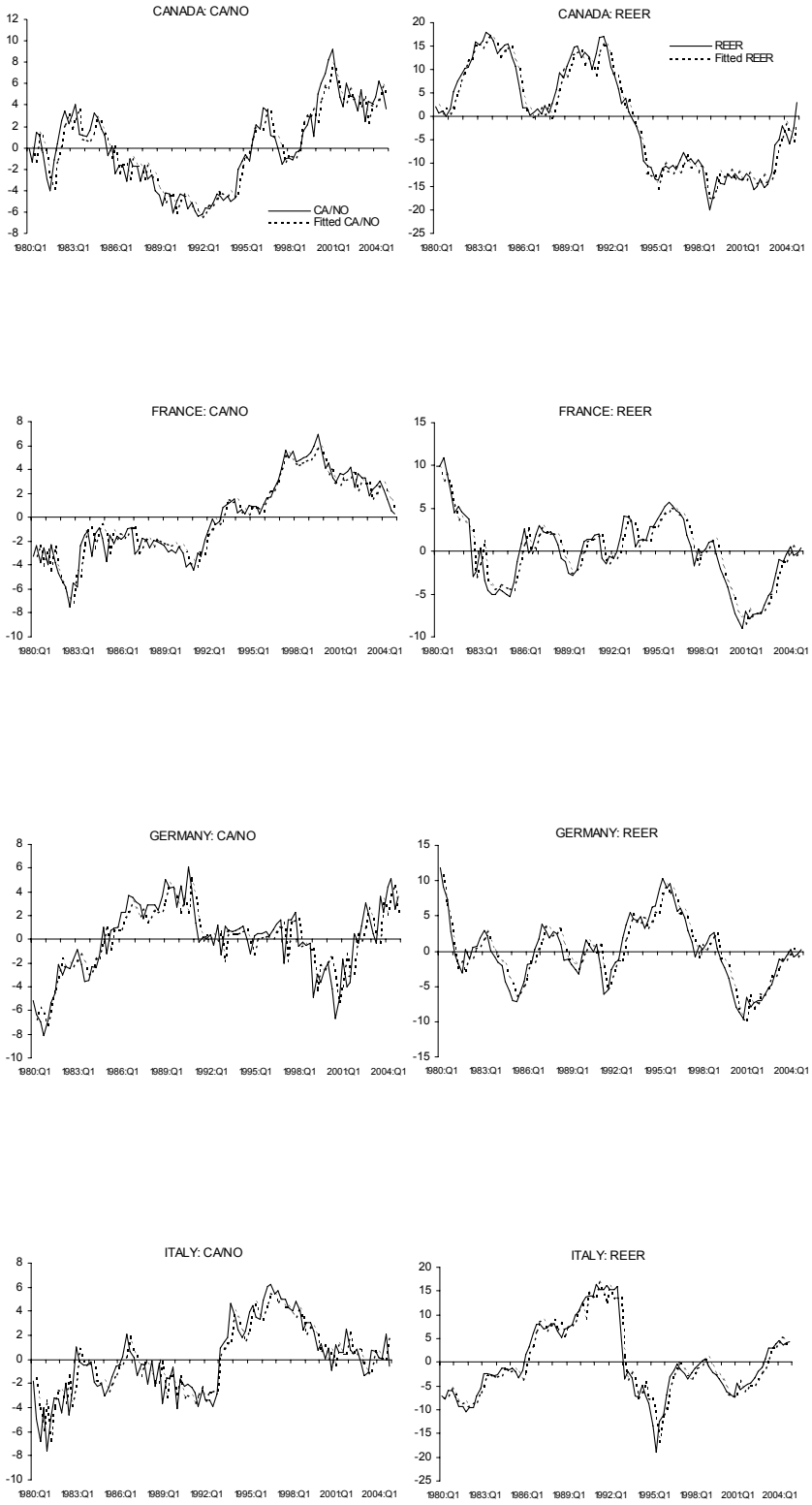
We think of this as potentially an useful nonlinear model for evaluating the hypothesis that threshold breaks in the real exchange rate have incremental predictive content for subsequent current account dynamics. Especially in a world of transport cost and trade frictions (Obstfeld and Rogoff, 2001) we should expect such threshold effects in real exchange rates to be potentially important in influencing trade flows. Given the nonlinear nature of the specification, we estimate the system as a set of simultaneous equations and conduct a three-stage grid search to identify the upper and lower reer threshold for each country.

Threshold system of de-meaned CA/NO and reer

Q1 1979- Q4 2004	Thresholds of REER (asymmetric band)			Slope coefficients CA/NO			Slope coefficients reer			Means	
COUNTRY	Upper threshold	Lower threshold		above	band	below	above	band	below	CA/NO	REER
CANADA	**10.67**	**-12.43**	**CA/NO{1}**	0.960***	0.927***	0.604***	0.110	0.251**	-0.182	**-1.519**	**473.31**
				(0.071)	(0.051)	(0.102)	(0.114)	(0.082)	(0.163)		
			reer{1}	-0.020	-0.012	-0.123***	0.943***	1.136***	0.870***		
				(0.019)	(0.022)	(0.033)	(0.030)	(0.035)	(0.052)		
					R^2: 0.891			*R^2: 0.966*			
FRANCE	**3.70**	**-1.28**	**CA/NO{1}**	1.202***	0.949***	0.940***	0.129	0.012	-0.123	**0.244**	**467.97**
				(0.071)	(0.041)	(0.048)	(0.108)	(0.062)	(0.073)		
			reer{1}	0.044	0.014	-0.020	0.863***	0.992***	0.907***		
				(0.039)	(0.054)	(0.034)	(0.059)	(0.082)	(0.051)		
					R^2: 0.925			*R^2: 0.889*			
GERMANY	**2.56**	**-2.02**	**CA/NO{1}**	0.875***	0.846***	0.814***	0.266***	-0.003	0.296**	Pre-/Post 1991	
				(0.091)	(0.070)	(0.118)	(0.090)	(0.069)	(0.116)	**6.576**	**473.76**
			reer{1}	-0.026	-0.206**	-0.017	0.903***	0.910***	0.866***	**3.536**	**468.46**
				(0.050)	(0.087)	(0.053)	(0.049)	(0.086)	(0.052)		
					R^2: 0.752			*R^2: 0.890*			
ITALY	**8.02**	**-6.25**	**CA/NO{1}**	-0.569*	0.872***	0.888***	1.583**	0.026	0.029	**-0.080**	**466.83**
				(0.343)	(0.063)	(0.075)	(0.526)	(0.097)	(0.115)		
			reer{1}	-0.300***	0.005	-0.039	1.240***	1.081***	0.892***		
				(0.070)	(0.037)	(0.032)	(0.108)	(0.057)	(0.049)		
					R^2: 0.797			*R^2: 0.927*			
JAPAN	**4.49**	**-13.28**	**CA/NO{1}**	0.804***	0.996***	0.862***	1.518**	1.599**	0.367	**4.289**	**438.15**
				(0.076)	(0.081)	(0.048)	(0.538)	(0.567)	(0.342)		
			reer{1}	-0.004	0.019	-0.007	0.932***	0.562***	0.934***		
				(0.005)	(0.017)	(0.005)	(0.036)	(0.124)	(0.036)		
					R^2: 0.906			*R^2: 0.951*			
UK	**8.05**	**-9.13**	**CA/NO{1}**	0.375**	0.728***	0.371*	-0.344*	-0.011	0.201	**-1.965**	**448.16**
				(0.149)	(0.089)	(0.208)	(0.207)	(0.124)	(0.290)		
			reer{1}	0.013	0.160**	-0.030	0.963***	1.112***	0.885***		
				(0.042)	(0.053)	(0.035)	(0.058)	(0.073)	(0.049)		
					R^2: 0.513			*R^2: 0.895*			
US	**3.48**	**-0.04**	**CA/NO{1}**	0.919***	1.202***	0.920***	0.531*	0.249	0.294	**-2.350**	**453.32**
				(0.029)	(0.048)	(0.038)	(0.193)	(0.319)	(0.254)		
			reer{1}	-0.018***	-0.040	-0.014**	0.997***	1.263***	0.988***		
				(0.005)	(0.031)	(0.007)	(0.033)	(0.206)	(0.048)		
					R^2: 0.968			*R^2: 0.943*			
					SE in brackets			*SE in brackets*			

Table 5.4: Threshold System of CA/NO and REER

Results are reported in Table 5.4. The country dynamics and thresholds slightly differ for some country from the univariate estimates. In general, the fit of the model is good and provides close predicted values (Figure 5.1 shows fitted versus actual values for the two variables for all G7 countries).


CANADA: CA/NO
CA/NO
Fitted CA/NO
CANADA: REER
REER
Fitted REER
FRANCE: CA/NO
FRANCE: REER
GERMANY: CA/NO
GERMANY: REER
ITALY: CA/NO
ITALY: REER
1980:Q1
1983:Q1
1986:Q1
1989:Q1
1992:Q1
1995:Q1
1998:Q1
2001Q1
2004:Q1

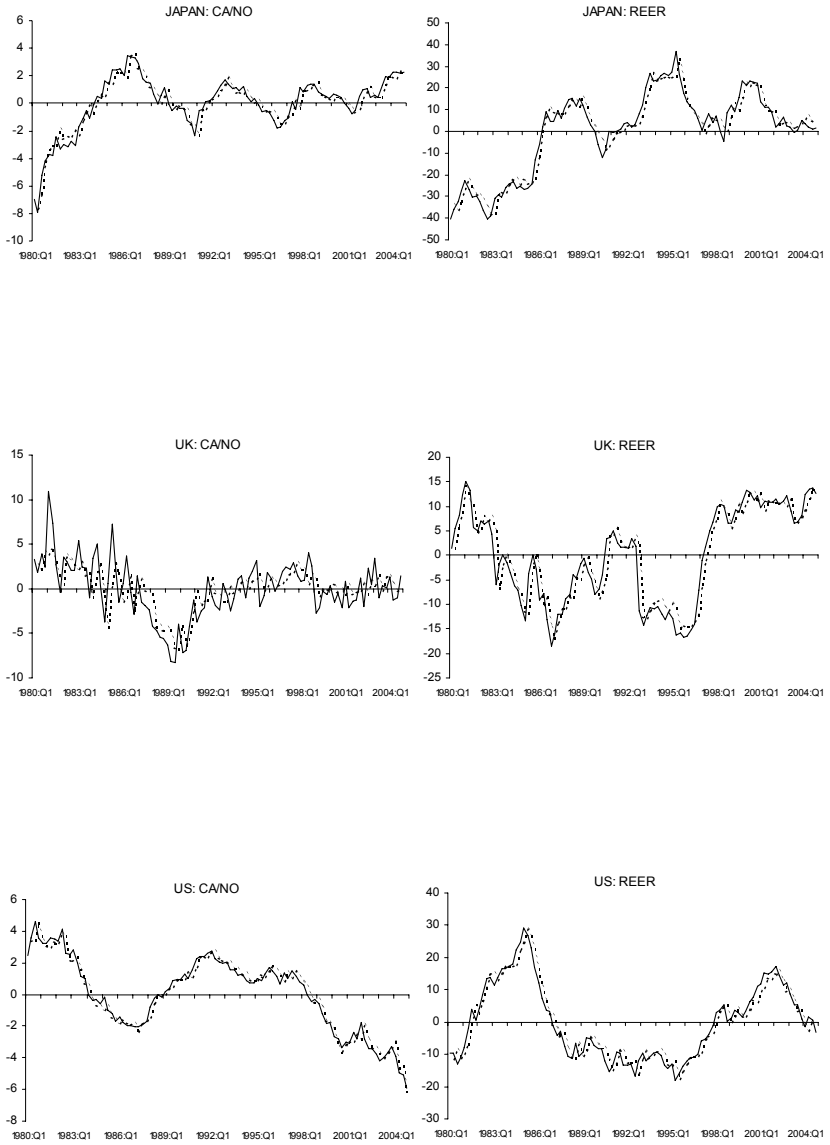


Figure 5.1: Threshold System of CA/NO and REER - Fitted versus Actual Values

5.5 Speed of CA and reer adjustment: nonlinear impulse responses

In order to better evaluate the speed of mean reversion exhibited by ca and q, we look at impulse responses, following the methodology introduced by Koop et al.(1996) and applied in Taylor et al.(2001). The difficulty in producing multi-step forecasts for a nonlinear process, given the impact of future disturbances on the nonlinear function, is solved by Monte Carlo integration. Impulse responses of ca and q to shocks of the variables of $\pm 5\%$, $\pm 10\%$ and $\pm 20\%$ are plotted for each country in Figures 5.2-8. The fact that the different impulse responses - with the exception of France, due to its persistent surplus slope coefficient[34] - converge to a single flat line profile is an indication that the processes are integrated short memory in mean, i.e. that the long-horizon prediction of the processes is a vector of constants - in this case, zeros as variables are demeaned. Figures 5.2-8 show different speed of adjustment across countries: the impulse responses of Germany and Japan converge to zero after around 40 quarters, while Canada and US only after over 70 quarters.

[34] French estimates suggest lack of mean-reversion during surplus periods. This specific pattern seems to arise from the ca-reer interaction: the French ca surplus kept increasing although its reer was appreciating in 1998 and then steadily declined despite the significant real depreciation in the 1999-2001 period. Therefore, the high surplus coefficient and the resulting explosive path are likely to reflect the lack of correction in the French ca -at least during surplus periods - following movements in the reer. Nevertheless, the ca behaviour remains sustainable once we simply look at the baseline model with thresholds of ca.

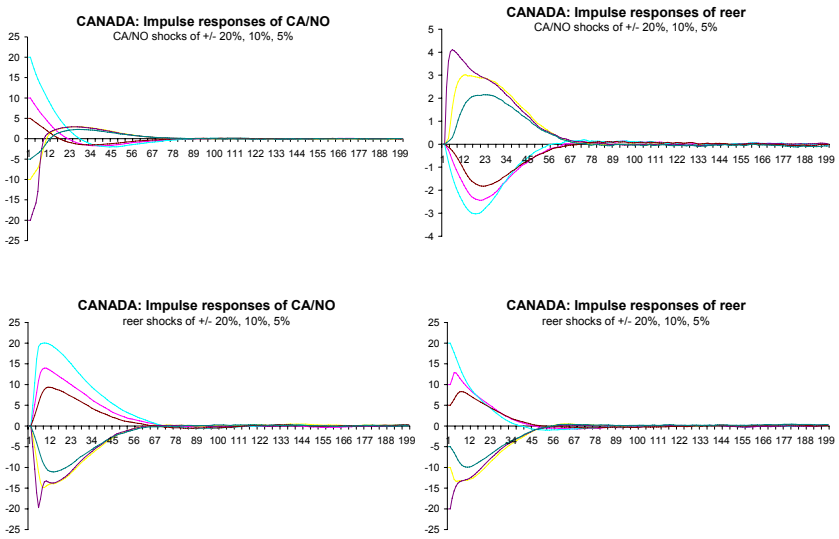


Figure 5.2: Nonlinear Impulse Responses - Canada

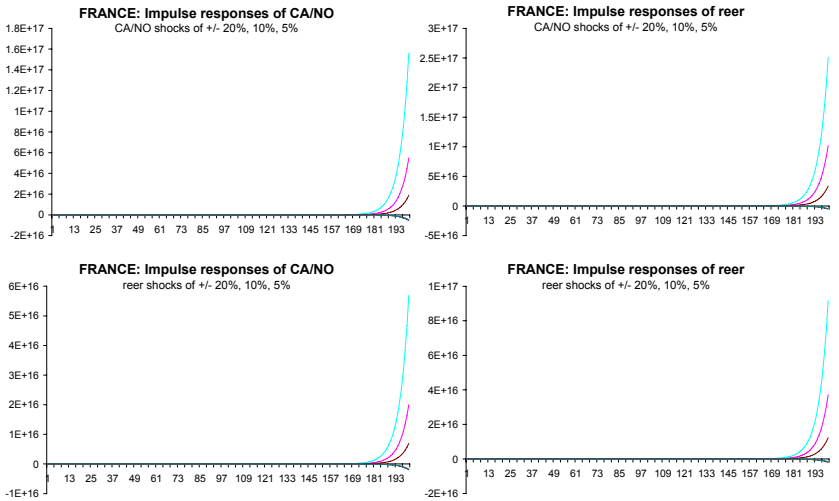


Figure 5.3: Nonlinear Impulse Responses - France

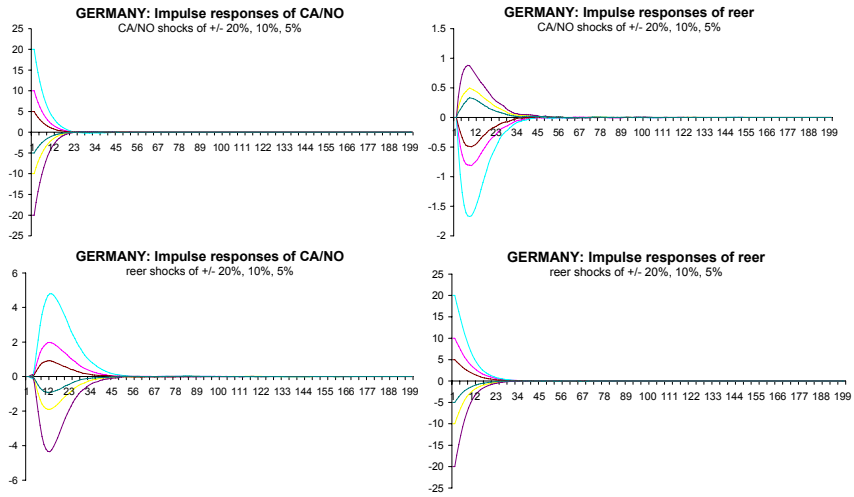


Figure 5.4: Nonlinear Impulse Responses - Germany

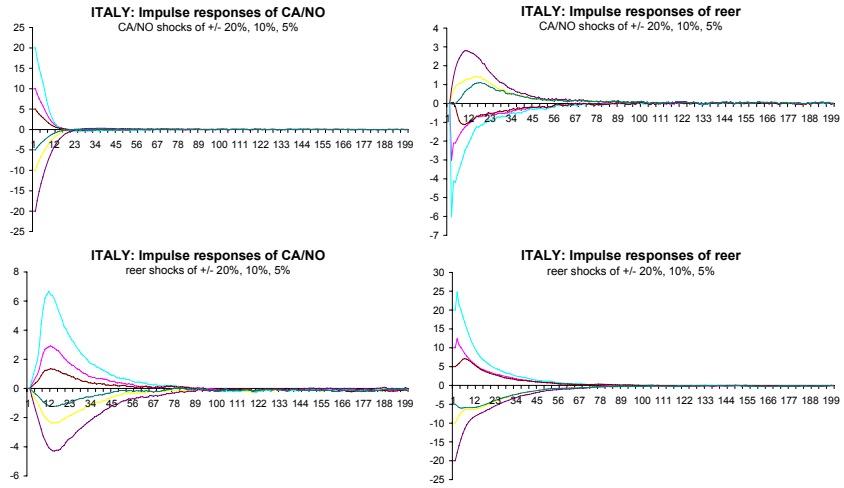


Figure 5.5: Nonlinear Impulse Responses - Italy

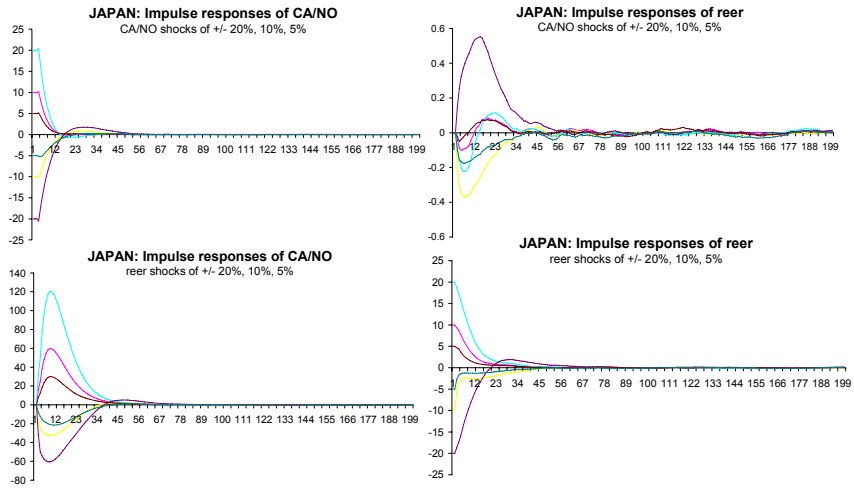


Figure 5.6: Nonlinear Impulse Responses - Japan

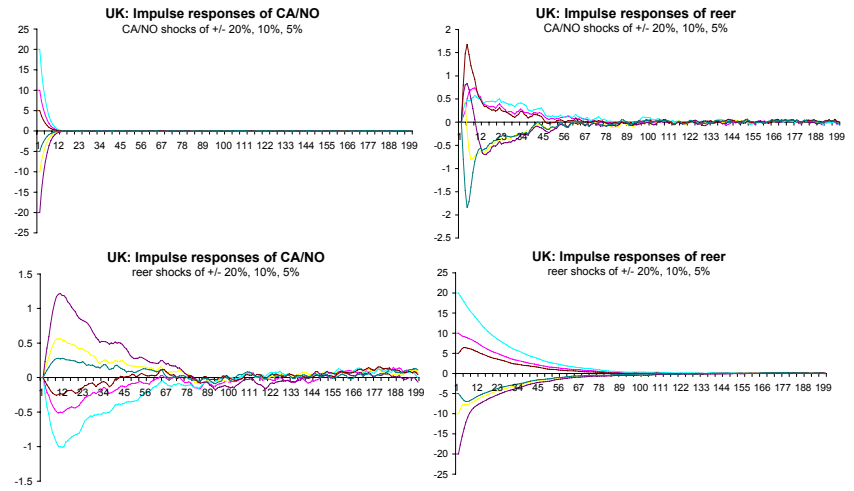


Figure 5.7: Nonlinear Impulse Responses - UK

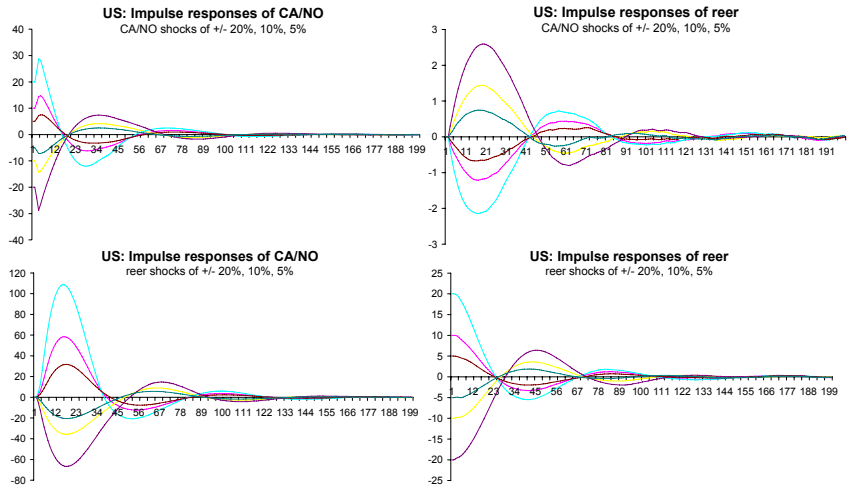


Figure 5.8: Nonlinear Impulse Responses - US

5.6 Concluding remarks

Are there thresholds of current account adjustment and are these identified by changes in the adjustment of the real exchange rate? This chapter has reported evidence in favour of both these propositions. We found statistically significant evidence of differing adjustment dynamics in the current-account-to-net-output ratio and real exchange rate for all of the G7 countries examined. In particular, each country displayed three regimes — a surplus regime and a deficit regime in which the current account tended to revert towards its long-run mean, albeit at different speeds in each regime (depending on the size of real exchange rate adjustment), and an 'inertia regime' in which, for intermediate levels of the real exchange rate misalignment, current account adjustment was negligible. We also showed significant cross-country variation in the size of the estimated thresholds and in the estimated speeds of adjustment once countries cross their relevant thresholds.

5.A Appendix: Data and variable definitions

We look at quarterly data for the period 1979:1 to 2004:4. We choose our starting date to begin six years after the advent of floating exchange rates and the initial globalization of the international capital market that occurred at that time and in conjunction with the first oil shock. The data in the analysis are obtained from the International Financial Statistics (IFS) Database by the IMF. All variables are seasonally adjusted and expressed in national currency. According to national account statistics, the current account variable is estimated as the sum of net exports and net primary income from abroad (NPIA); net output is obtained by subtracting Government consumption expenditure and gross fixed capital formation (investment) to GDP. For the exchange rate variables, we refer to trade weighted real effective exchange rate indices, whereas a decrease in the index corresponds to a depreciation (data are available as from 1980:1).

Chapter 6
A TAR test of the present-value model of the current account

6.1 Introduction

The intertemporal approach to the current account provides an intuitive explanation to the existence of temporary current account deficits and surpluses, by focusing on the optimal saving decisions of a representative household as it smooths consumption over time. In its simplest form, the model implies that a country current account position is determined by the present value of expected future changes in net output, such that a country experiencing a temporary fall in output would be expected to run a current account deficit and viceversa.

Intertemporal models of the current account have been estimated and tested many times, and their high frequency implications – that current account dynamics are fully described by the discounted sum of future changes in net output – are usually rejected. Most of these empirical works have extended the seminal work by Campbell (1987) and Campbell-Shiller (1987) on consumption theory, by comparing VAR forecasts of the present value of changes in net output to the actual current account series. Results proved unsatisfactory, especially for small open economies. Nevertheless, contributions by Bergin-Sheffrin(2000) and recently by Nason-Rogers (2006) have highlighted the relevant explanatory power of the interest rate and real exchange rate in the present value model of the cur-

rent account. Although the restrictions of the model are still rejected for some countries, the inclusion of such valuation effects significantly improves the model predictions.

In this chapter, we build on the results of chapter 5 in order to estimate a present value test of the current account within a nonlinear framework. Evidence on the threshold behaviour of the current account and on the significant effect of the real exchange rate on its adjustment encourage this type of analysis. The next section of the chapter sketches the main features of the theoretical model used to develop our econometric framework. Section 3 describes the empirical methodology and presents the estimation results for the nonlinear system as well as the parameter values for the present value test. Section 4 analyses the temporary components of the relevant variables in order to identify their convergence patterns towards steady state and therefore enable a finite time analysis of the model. The predictive power of the model is finally analysed in Section 5. Section 6 concludes the chapter.

6.2 The present value model of the CA: theoretical framework

The workhorse intertemporal model of the current account (Sachs, 1981; Sheffrin and Woo, 1990, via Campbell, 1987) has the following strong empirical implication:

$$CA_t = -E_t \sum_{i=1}^{\infty} (1+r)^{-i} \Delta NO_{t+i} \tag{6.28}$$

Thus, according to the intertemporal model, the current account balance is the discounted present value of future expected changes in net output. Papers that implement tests of this equation have traditionally assumed that the level of NO_t is $I(1)$ so that the first

difference of NO_t, ΔNO_t is $I(0)$. This assumption has a very important implication: that the level of the current account is stationary. This equation has been tested many times in a linear vector autoregression framework and is almost always rejected.

For a variety of reasons, it would seem to be preferable to model $\Delta \ln NO_t = \Delta no_t$ as stationary. Following Campbell and Shiller (1987), Kano (2003) proposes the following present value representation of the current account-net output ratio:

$$\begin{aligned} \frac{CA_t}{NO_t} &= br_t + [(\sigma - 1)c + 1]\sum_{i=1}^{\infty} k^i E_t r_{t+i} - \sum_{i=1}^{\infty} k^i E_t \Delta \ln NO_{t+i} \qquad (6.29) \\ k &= e^{\gamma - \mu} \end{aligned}$$

where $c, b, \gamma, \mu, \sigma$ are respectively the unconditional mean of the current account-net output ratio, $\frac{CA_t}{NO_t}$; the unconditional mean of the net foreign asset-net output ratio, $\frac{B_t}{NO_{t-1}}$; the unconditional mean of the first difference of the log of net output, $\Delta \ln NO_t$; the unconditional mean of the gross world real interest rate, $\ln(1 + r_t) \approx r_t$; the elasticity of intertemporal substitution. All the variables are represented as deviations from their unconditional mean. Note that if the log difference of net output is stationary, it is the current account to net output ratio which is stationary, not simply the current account itself.

In particular, the statistically significant effect of the real exchange rate on the dynamic equation for the current-account-to-net-output-ratio suggests the relevance of the introduction of this variable in the intertemporal model. Along this line, Bergin-Sheffrin (2000) consider a small country producing traded and nontraded goods and introduce a

consumption-based real interest rate r^*, which reflects both the interest rate r and the change in the relative price of nontraded goods, $reer$.

$$r_t^* = r_t + [\frac{1-\sigma}{\sigma}(1-a)]\Delta q_t \tag{6.30}$$

where

$$U(C_{Tt}, C_{Nt}) = \frac{1}{1-1/\sigma}(C_{Tt}^a C_{Nt}^{1-a})^{1-1/\sigma} \tag{6.31}$$

$$\Delta q_t = \ln REER_t - \ln REER_{t-1} \tag{6.32}$$

Equations (6.29) and (6.30) suggest the estimation of the following present-value representation:

$$\frac{CA_t}{NO_t} = br_t^* + [(\sigma-1)c+1]\sum_{i=1}^{\infty} k^i E_t r_{t+i}^* - \sum_{i=1}^{\infty} k^i E_t \Delta \ln NO_{t+i} \tag{6.33}$$

or alternatively,

$$\begin{aligned} \frac{CA_t}{NO_t} &= br_t + b[\frac{1-\sigma}{\sigma}(1-a)]\Delta reer_t + \\ &+[(\sigma-1)c+1]\sum_{i=1}^{\infty} k^i E_t[r_{t+i} + [\frac{1-\sigma}{\sigma}(1-a)]\Delta reer_{t+i}] + \\ &-\sum_{i=1}^{\infty} k^i E_t \Delta \ln NO_{t+i} \end{aligned} \tag{6.34}$$

This new framework for testing the present value model of the current account allows to account not just for expected changes in net output, but also for both intertemporal and intratemporal changes in the exchange rate and the interest rate. Under this setting, consumption decisions are affected by valuation effects arising from changes in the terms

of borrowing and lending with the rest of the world as well as from changes in the relative price of nontraded goods.

6.3 Empirical methodology and parameter calibration

In this section of the chapter, we plan to estimate a system of ca_t, q_t, Δno_t and r_t, whereas ca_t and q_t follow a nonlinear threshold process with thresholds of q_t[35]. Note that the use of a nonlinear framework for the analysis of the intertemporal approach to the current account does not allow to test the standard restrictions of the PVM as in a linear VAR model. Therefore, we need to compute a finite-time present-value model prediction for the ca variable directly from our forecasts of the variables, obtained by bootstrap methods on the system fitted residuals.

The standard VAR model used in the literature is replaced in our analysis by a group of simultaneous equation with the following processes:

$$\begin{aligned} ca_t &= \overline{\alpha}_1 \times 1\{q_{t-d}, \overline{\delta}\} \times ca_{t-1} + \underline{\alpha}_2 \times 1\{q_{t-d}, \underline{\delta}\} \times ca_{t-1} + \\ &+ \overline{\underline{\alpha}}_3 \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times ca_{t-1} + \\ &+ \overline{\beta}_1 \times 1\{q_{t-d}, \overline{\delta}\} \times q_{t-1} + \underline{\beta}_2 \times 1\{q_{t-d}, \underline{\delta}\} \times q_{t-1} + \\ &+ \overline{\underline{\beta}}_3 \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times q_{t-1} + \\ &+ \gamma_1 \times \Delta no_{t-1} + \delta_1 \times r_t^* + u_{1t} \end{aligned} \tag{6.35}$$

[35] We reject the hypothesis of nonlinearities in the behaviour of both Δno_t and r_t^*.

$$
\begin{aligned}
q_t &= \overline{\alpha}_4 \times 1\{q_{t-d}, \overline{\delta}\} \times ca_{t-1} + \underline{\alpha}_5 \times 1\{q_{t-d}, \underline{\delta}\} \times ca_{t-1} + \\
&+ \underline{\overline{\alpha}}_6 \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times ca_{t-1} + \\
&+ \overline{\beta}_4 \times 1\{q_{t-d}, \overline{\delta}\} \times q_{t-1} + \underline{\beta}_5 \times 1\{q_{t-d}, \underline{\delta}\} \times q_{t-1} + \\
&+ \underline{\overline{\boldsymbol{\beta}}}_6 \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times q_{t-1} + \\
&+ \gamma_2 \times \Delta no_{t-1} + \delta_2 \times r_t^* + u_{2t} \qquad (6.36)
\end{aligned}
$$

$$
\begin{aligned}
\Delta no_t &= \overline{\alpha}_7 \times 1\{q_{t-d}, \overline{\delta}\} \times ca_{t-1} + \underline{\alpha}_8 \times 1\{q_{t-d}, \underline{\delta}\} \times ca_{t-1} + \\
&+ \underline{\overline{\alpha}}_9 \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times ca_{t-1} + \\
&+ \overline{\beta}_7 \times 1\{q_{t-d}, \overline{\delta}\} \times q_{t-1} + \underline{\beta}_8 \times 1\{q_{t-d}, \underline{\delta}\} \times q_{t-1} + \\
&+ \underline{\overline{\boldsymbol{\beta}}}_9 \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times q_{t-1} + \\
&+ \gamma_3 \times \Delta no_{t-1} + \delta_3 \times r_t^* + u_{3t} \qquad (6.37)
\end{aligned}
$$

$$
\begin{aligned}
r_t^* &= \overline{\alpha}_{10} \times 1\{q_{t-d}, \overline{\delta}\} \times ca_{t-1} + \underline{\alpha}_{11} \times 1\{q_{t-d}, \underline{\delta}\} \times ca_{t-1} + \\
&+ \underline{\overline{\alpha}}_{12} \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times ca_{t-1} + \\
&+ \overline{\beta}_{10} \times 1\{q_{t-d}, \overline{\delta}\} \times q_{t-1} + \underline{\beta}_{11} \times 1\{q_{t-d}, \underline{\delta}\} \times q_{t-1} + \\
&+ \underline{\overline{\boldsymbol{\beta}}}_{12} \times (1 - 1\{q_{t-d}, \overline{\delta}\}) \times (1 - 1\{q_{t-d}, \underline{\delta}\}) \times q_{t-1} + \\
&+ \gamma_4 \times \Delta no_{t-1} + \delta_4 \times r_t^* + u_{4t} \qquad (6.38)
\end{aligned}
$$

Estimates of the system are reported in Table 6.1-4 and fitted values for the relevant variables are plotted in Figure 6.1-7. Data sources and definitions are described in the Appendix. Results indicate a good fit for the ca, q and r^*. equations in all G7 countries. However, the Dno equations present unsatisfactory R^2. In general. the system does not seem able to capture the volatility of changes in net output in line with actual data. This

result represents a strong limitation to any test of a present-value model of the current account, given the essential role of Dno forecasts for the ca predictions.

CANADA

	Dno						CA/NO					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**0.092**	0.091	**0.092**	0.091	**0.092**	0.091	**-0.086**	0.106	**-0.086**	0.106	**-0.086**	0.106
WRIR{1}	**-0.219**	0.095	**-0.219**	0.095	**-0.219**	0.095	**0.050**	0.111	**0.050**	0.111	**0.050**	0.111
CA/NO{1}	**0.235**	0.060	**-0.095**	0.045	**-0.160**	0.087	**0.955**	0.069	**0.940**	0.053	**0.603**	0.101
reer{1}	**0.064**	0.020	**-0.035**	0.020	**-0.018**	0.029	**-0.023**	0.023	**-0.021**	0.024	**-0.133**	0.034
R^2	0.257						R^2	0.898				

	reer						WRIR					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**0.093**	0.175	**0.093**	0.175	**0.093**	0.175	**0.050**	0.053	**0.050**	0.053	**0.050**	0.053
WRIR{1}	**-0.073**	0.183	**-0.073**	0.183	**-0.073**	0.183	**0.735**	0.055	**0.735**	0.055	**0.735**	0.055
CA/NO{1}	**0.093**	0.114	**0.244**	0.087	**-0.197**	0.167	**0.069**	0.035	**-0.025**	0.026	**-0.093**	0.051
reer{1}	**0.954**	0.038	**1.136**	0.039	**0.864**	0.056	**0.036**	0.011	**0.016**	0.012	**0.015**	0.017
R^2	0.966						R^2	0.853				

Upper threshold: 10.44 Lower threshold: -13.04

FRANCE

	Dno						CA/NO					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.715**	0.068	**-0.715**	0.068	**-0.715**	0.068	**-0.065**	0.012	**-0.065**	0.012	**-0.065**	0.012
WRIR{1}	**0.237**	0.378	**0.237**	0.378	**0.237**	0.378	**-0.187**	0.069	**-0.187**	0.069	**-0.187**	0.069
CA/NO{1}	**0.135**	0.230	**-0.254**	0.418	**0.106**	0.277	**0.983**	0.042	**1.022**	0.076	**0.760**	0.050
reer{1}	**-0.115**	0.181	**0.612**	0.417	**-0.016**	0.171	**0.004**	0.033	**-0.048**	0.076	**-0.038**	0.031
R^2	0.558						R^2	0.941				

	reer						WRIR					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**0.016**	0.021	**0.016**	0.021	**0.016**	0.021	**0.010**	0.009	**0.010**	0.009	**0.010**	0.009
WRIR{1}	**-0.174**	0.116	**-0.174**	0.116	**-0.174**	0.116	**0.871**	0.051	**0.871**	0.051	**0.871**	0.051
CA/NO{1}	**0.014**	0.071	**-0.032**	0.129	**-0.180**	0.085	**-0.020**	0.031	**-0.028**	0.057	**-0.071**	0.037
reer{1}	**0.816**	0.056	**0.875**	0.128	**0.899**	0.053	**0.004**	0.024	**0.241**	0.056	**0.013**	0.023
R^2	0.877						R^2	0.868				

Upper threshold: 0.60 Lower threshold: -0.20

Table 6.1: Threshold Model for Canada and France

GERMANY

	Dno						CA/NO					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.114**	0.095	**-0.114**	0.095	**-0.114**	0.095	**-0.067**	0.063	**-0.067**	0.063	**-0.067**	0.063
WRIR{1}	**0.274**	0.143	**0.274**	0.143	**0.274**	0.143	**0.017**	0.096	**0.017**	0.096	**0.017**	0.096
CA/NO{1}	**0.089**	0.136	**0.741**	0.204	**-0.005**	0.118	**0.840**	0.091	**0.759**	0.136	**0.852**	0.079
reer{1}	**-0.048**	0.079	**0.853**	0.273	**0.013**	0.078	**-0.003**	0.052	**-0.258**	0.182	**-0.053**	0.052
R^2	0.168						R^2	0.728				

	reer						WRIR					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.207**	0.061	**-0.207**	0.061	**-0.207**	0.061	**-0.021**	0.025	**-0.021**	0.025	**-0.021**	0.025
WRIR{1}	**0.031**	0.093	**0.031**	0.093	**0.031**	0.093	**0.913**	0.037	**0.913**	0.037	**0.913**	0.037
CA/NO{1}	**0.164**	0.088	**0.055**	0.132	**0.137**	0.076	**-0.098**	0.035	**-0.052**	0.053	**-0.011**	0.031
reer{1}	**0.894**	0.051	**0.629**	0.177	**0.904**	0.051	**-0.002**	0.020	**0.261**	0.071	**0.037**	0.020
R^2	0.886						R^2	0.870				

Upper threshold: 1.22 Lower threshold: -0.01

ITALY

	Dno						CA/NO					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.136**	0.103	**-0.136**	0.103	**-0.136**	0.103	**-0.019**	0.130	**-0.019**	0.130	**-0.019**	0.130
WRIR{1}	**0.129**	0.068	**0.129**	0.068	**0.129**	0.068	**-0.061**	0.086	**-0.061**	0.086	**-0.061**	0.086
CA/NO{1}	**-0.097**	0.253	**0.006**	0.099	**0.026**	0.043	**-0.538**	0.320	**0.594**	0.125	**0.873**	0.054
reer{1}	**-0.039**	0.054	**0.018**	0.038	**0.003**	0.021	**-0.304**	0.068	**-0.060**	0.048	**-0.028**	0.026
R^2	0.056						R^2	0.815				

	reer						WRIR					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.197**	0.218	**-0.197**	0.218	**-0.197**	0.218	**0.090**	0.065	**0.090**	0.065	**0.090**	0.065
WRIR{1}	**0.116**	0.145	**0.116**	0.145	**0.116**	0.145	**0.877**	0.043	**0.877**	0.043	**0.877**	0.043
CA/NO{1}	**1.609**	0.538	**-0.100**	0.210	**0.071**	0.092	**-0.030**	0.159	**-0.245**	0.062	**-0.045**	0.027
reer{1}	**1.251**	0.115	**1.070**	0.080	**0.945**	0.044	**-0.011**	0.034	**-0.030**	0.024	**-0.006**	0.013
R^2	0.925						R^2	0.853				

Upper threshold: 7.83 Lower threshold: 0.00

Table 6.2: Threshold model estimations for Germany and Italy

JAPAN

	Dno						**CA/NO**					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.275**	0.100	**-0.275**	0.100	**-0.275**	0.100	**-0.088**	0.057	**-0.088**	0.057	**-0.088**	0.057
WRIR{1}	**0.018**	0.069	**0.018**	0.069	**0.018**	0.069	**-0.138**	0.039	**-0.138**	0.039	**-0.138**	0.039
CA/NO{1}	**0.030**	0.141	**-0.030**	0.126	**0.044**	0.094	**0.879**	0.081	**0.755**	0.072	**1.052**	0.054
reer{1}	**-0.011**	0.008	**-0.034**	0.031	**-0.014**	0.010	**-0.011**	0.005	**0.027**	0.018	**-0.022**	0.005
R^2	0.115						R^2	0.901				

	reer						**WRIR**					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**0.288**	0.450	**0.288**	0.450	**0.288**	0.450	**0.034**	0.065	**0.034**	0.065	**0.034**	0.065
WRIR{1}	**-0.120**	0.309	**-0.120**	0.309	**-0.120**	0.309	**0.868**	0.045	**0.868**	0.045	**0.868**	0.045
CA/NO{1}	**1.873**	0.636	**1.208**	0.566	**0.910**	0.422	**0.152**	0.092	**-0.046**	0.082	**-0.075**	0.061
reer{1}	**0.943**	0.038	**0.654**	0.139	**0.914**	0.043	**-0.003**	0.006	**-0.058**	0.020	**-0.006**	0.006
R^2	0.949						R^2	0.868				

Upper threshold: 5.14 Lower threshold: -9.28

UK

	Dno						**CA/NO**					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.068**	0.102	**-0.068**	0.102	**-0.068**	0.102	**0.052**	0.047	**0.052**	0.047	**0.052**	0.047
WRIR{1}	**0.080**	0.356	**0.080**	0.356	**0.080**	0.356	**-0.076**	0.163	**-0.076**	0.163	**-0.076**	0.163
CA/NO{1}	**-0.043**	0.362	**-0.316**	0.302	**0.052**	0.233	**0.402**	0.165	**0.858**	0.138	**0.634**	0.106
reer{1}	**0.025**	0.118	**0.048**	0.127	**-0.019**	0.074	**0.005**	0.054	**0.105**	0.058	**-0.002**	0.034
R^2	0.020						R^2	0.486				

	reer						**WRIR**					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.116**	0.060	**-0.116**	0.060	**-0.116**	0.060	**-0.009**	0.013	**-0.009**	0.013	**-0.009**	0.013
WRIR{1}	**-0.378**	0.211	**-0.378**	0.211	**-0.378**	0.211	**0.864**	0.044	**0.864**	0.044	**0.864**	0.044
CA/NO{1}	**-0.255**	0.214	**0.356**	0.178	**-0.169**	0.137	**0.089**	0.045	**0.117**	0.037	**-0.028**	0.029
reer{1}	**0.884**	0.070	**1.045**	0.075	**0.914**	0.044	**-0.030**	0.015	**0.005**	0.016	**0.001**	0.009
R^2	0.905						R^2	0.849				

Upper threshold: 10.05 Lower threshold: -1.74

Table 6.3: Threshold model estimations for Japan and UK

US

	Dno						**CA/NO**					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.158**	0.102	**-0.158**	0.102	**-0.158**	0.102	**-0.215**	0.049	**-0.215**	0.049	**-0.215**	0.049
WRIR{1}	**-0.074**	0.075	**-0.074**	0.075	**-0.074**	0.075	**0.101**	0.036	**0.101**	0.036	**0.101**	0.036
CA/NO{1}	**0.380**	0.123	**-0.162**	0.058	**0.339**	0.216	**0.853**	0.059	**0.904**	0.028	**0.885**	0.104
reer{1}	**0.036**	0.013	**-0.036**	0.016	**0.041**	0.025	**-0.022**	0.006	**-0.029**	0.008	**-0.010**	0.012
R^2	0.242						R^2	0.968				

	reer						**WRIR**					
	above		band		below		above		band		below	
	Coeff	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*	*Coeff*	*SE*
Dno{1}	**-0.247**	0.368	**-0.247**	0.368	**-0.247**	0.368	**-0.097**	0.078	**-0.097**	0.078	**-0.097**	0.078
WRIR{1}	**-0.344**	0.271	**-0.344**	0.271	**-0.344**	0.271	**0.743**	0.057	**0.743**	0.057	**0.743**	0.057
CA/NO{1}	**0.933**	0.443	**0.320**	0.209	**0.521**	0.778	**0.285**	0.094	**0.091**	0.044	**0.036**	0.165
reer{1}	**1.034**	0.046	**1.068**	0.057	**0.984**	0.091	**0.028**	0.010	**-0.026**	0.012	**0.027**	0.019
R^2	0.944						R^2	0.882				

Upper threshold: 11.72 Lower threshold: -10.35

Table 6.4: Threshold model estimations for the US

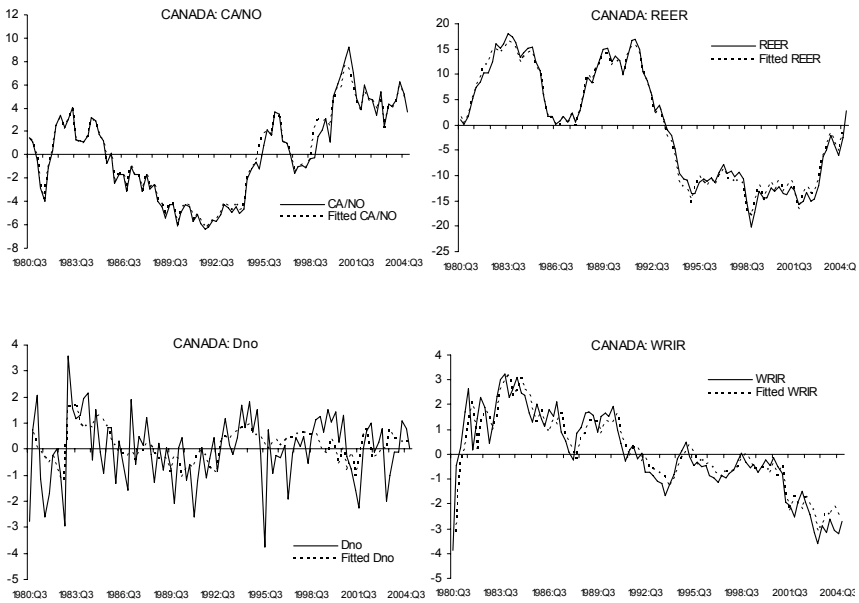


Figure 6.1: Canada - Fitted versus actual values

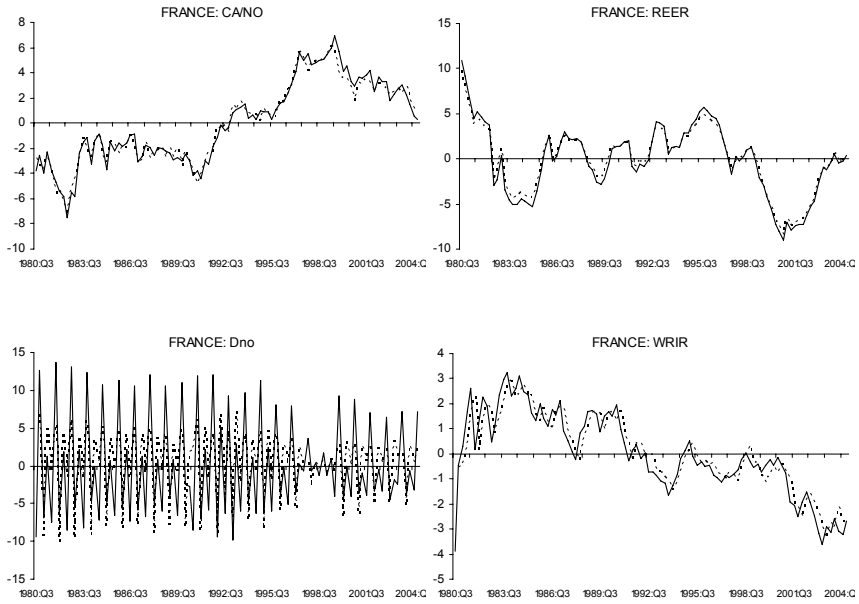


Figure 6.2: France - Fitted versus actual values

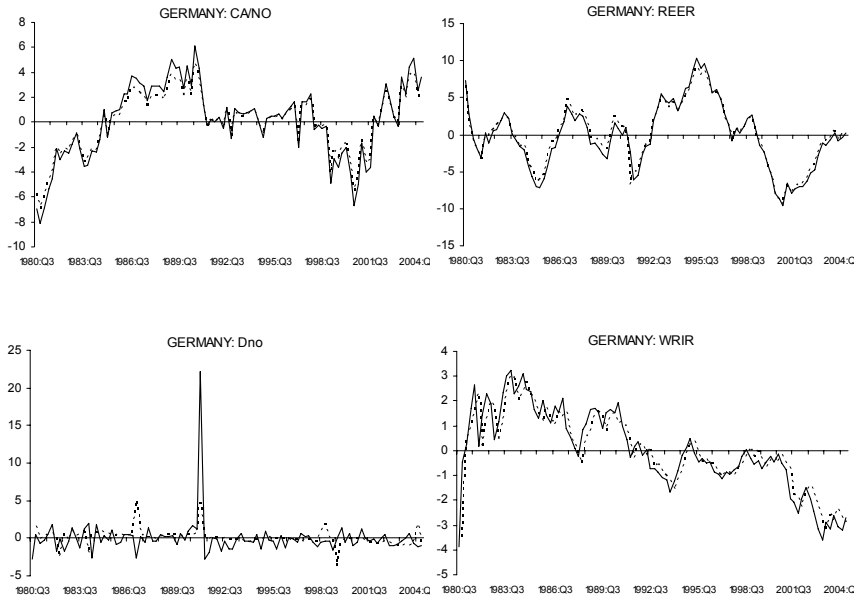


Figure 6.3: Germany - Fitted versus actual values

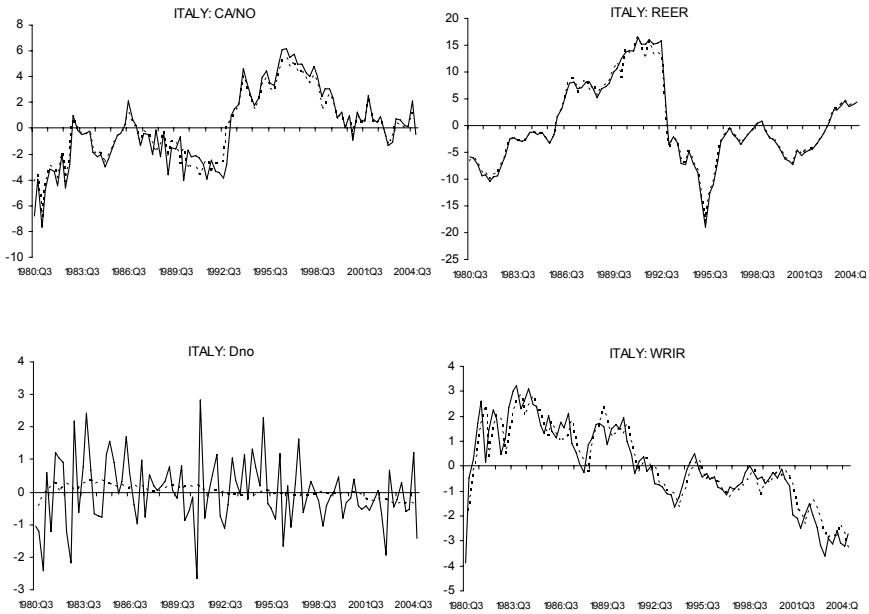


Figure 6.4: Italy - Fitted versus actual values

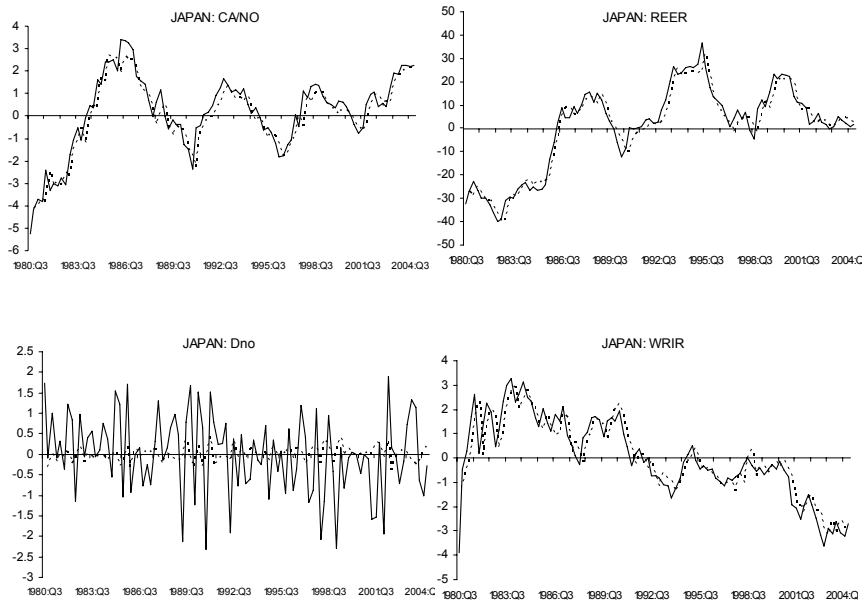


Figure 6.5: Japan - Fitted versus actual values

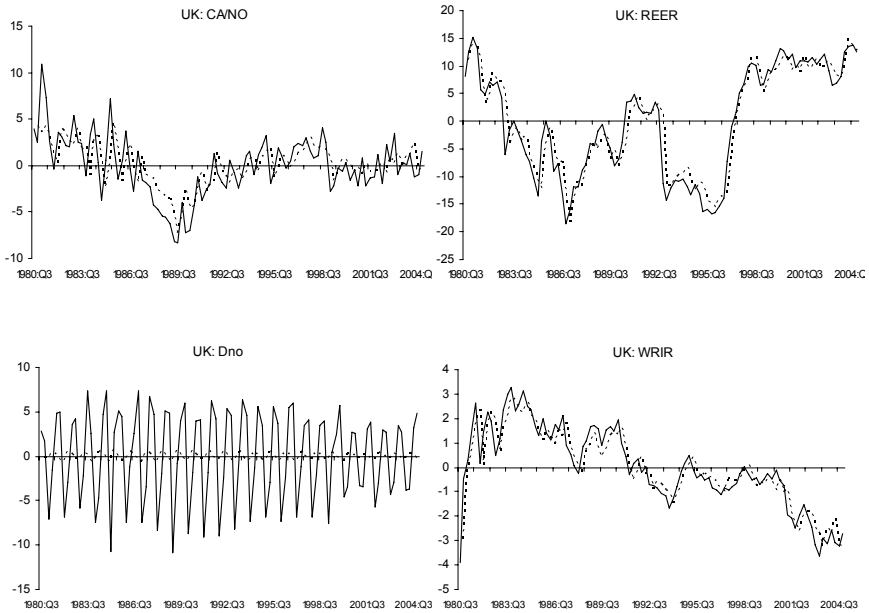


Figure 6.6: United Kingdom - Fitted versus actual values

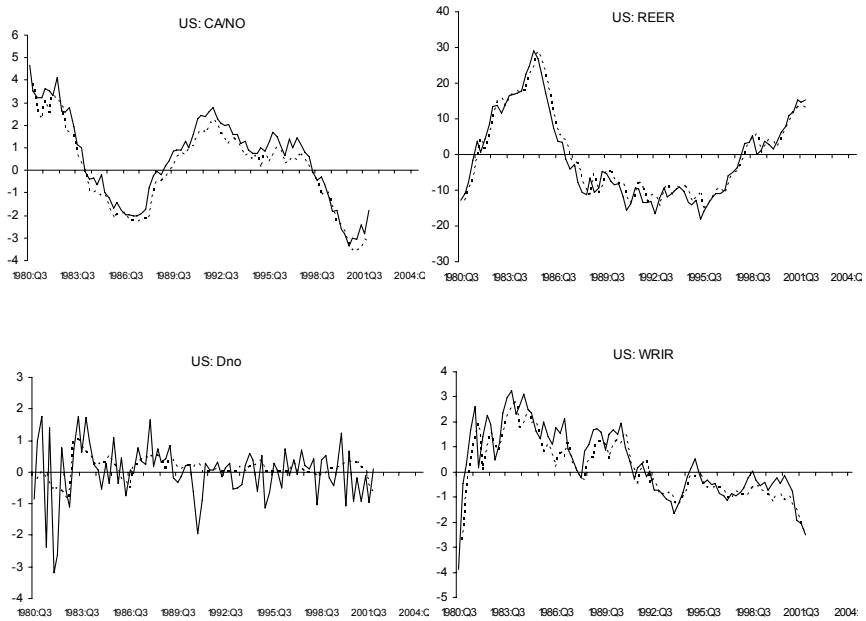


Figure 6.7: United States - Fitted versus actual values

In order to proceed with a test of the present-value condition (6.33), we first need to calibrate appropriate values of the parameters a, b, c, k (i.e. of γ, μ and η) and σ, according to the definitions in the previous section. Table 6.5 lists the parameters choices for each country. We choose a value of 0.5 for the traded goods share in private final consumption, a, in line with estimates by Stockman and Tesar (1995). The elasticity of intertemporal substitution, σ, is chosen for each country to minimise the mean squared error between the PVM predictions on the current account-net output ratio and its actual values. Nevertheless, we select values of σ in between 0.1 and 0.55 as generally suggested in the previous literature.[36]

[36] See Hall(1988) and Mehra and Prescott (1985).

Calibrated Parameters

	a	b	c	k	
Canada	0.50	-0.568	0.407	0.983	0.20
France	0.50	0.355	0.394	0.981	0.17
Germany	0.50	0.673	0.511	0.982	0.15
Italy	0.50	0.168	0.266	0.986	0.28
Japan	0.50	-0.280	0.570	0.980	0.48
UK	0.50	0.274	0.354	0.982	0.55
US	0.50	-0.073	0.443	0.981	0.11

Table 6.5: List of Calibrated Parameters

6.4 Speeds of adjustment: profile bundles

In order to define the length of the variable adjustments towards steady state and thus be able to identify a finite time window for our analysis, we look at the permanent and temporary components of the variables. We refer to the nonlinear methodology by Clarida and Taylor (2003). The authors apply the Beveridge-Nelson decomposition to nonlinear processes. The difficulty in producing multi-step forecasts for a nonlinear process, given the impact of future disturbances on the nonlinear function, is solved by Monte Carlo integration. We consider all possible combinations of starting values for the system variables of $\pm 5\%$ and their median values and then plot for each country the resulting profile bundles of ca, q, Dno and r^* in Figure 6.8-14 [37]. The different profile bundles all tend to converge to a flat line. This fact is an indication that the long-horizon prediction of the processes

[37] Due to space constraints, we report only results for these three shocks. Nevertheless, we obtain qualitatively equivalent results for shocks of different sizes.

is a vector of constants[38]. In general, an analysis of the profile bundles suggest that most countries converge to a finite value after 100 replications.

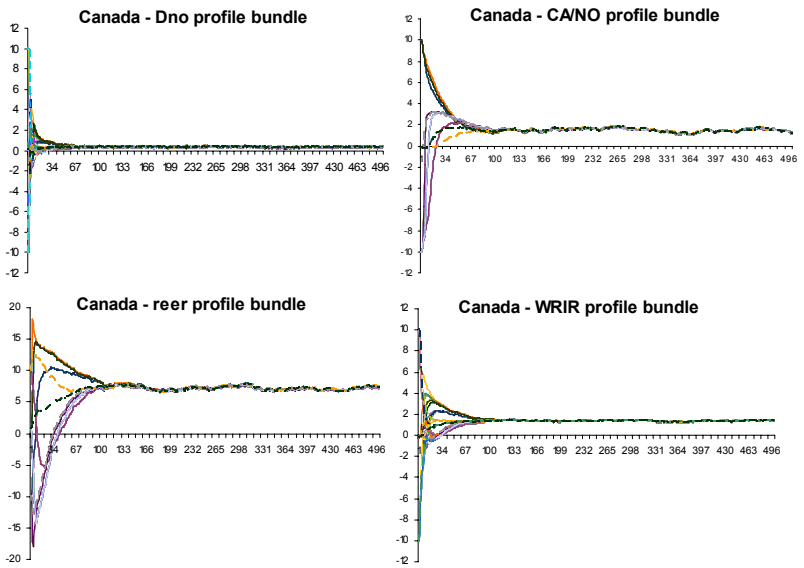


Figure 6.8: Canada Profile Bundles

[38] Note that, although variables are demeaned, their long run value is not necessarily zero given that the steady state mean does not necessarily coincide with the sample one.

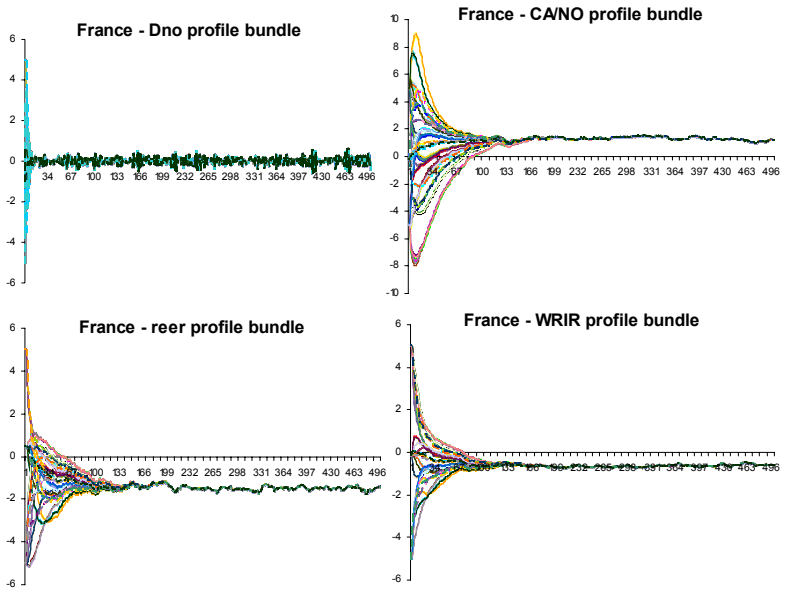


Figure 6.9: France Profile Bundles

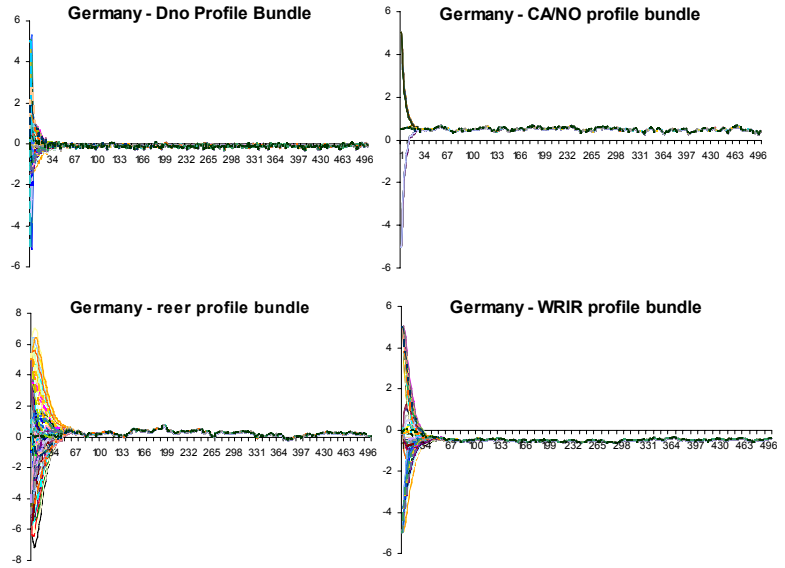


Figure 6.10: Germany Profile Bundles

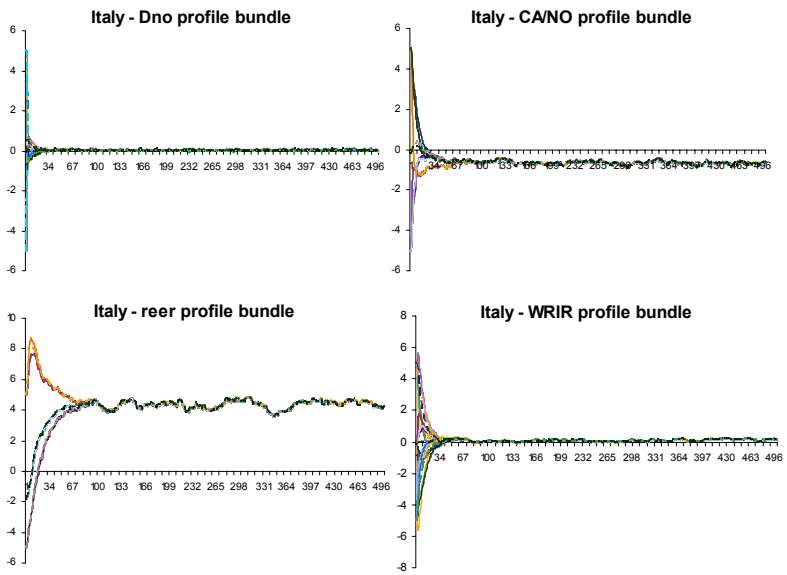


Figure 6.11: Italy Profile Bundles

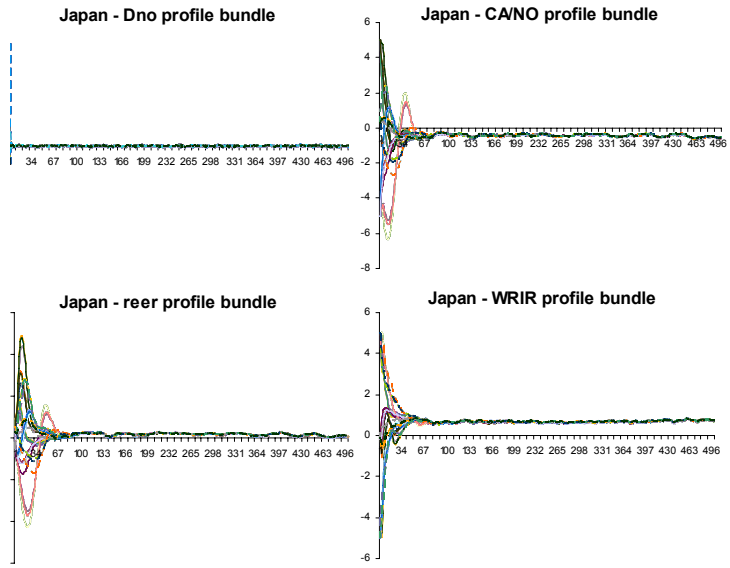


Figure 6.12: Japan Profile Bundles

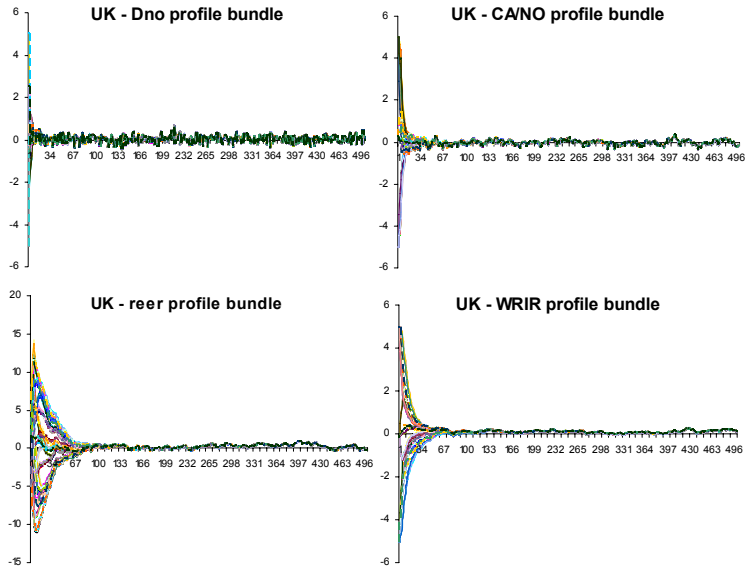


Figure 6.13: UK Profile Bundles

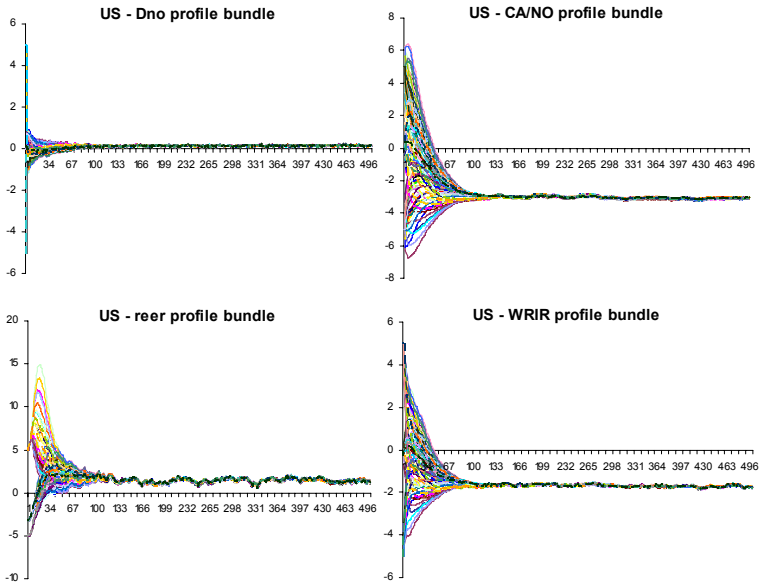


Figure 6.14: US Profile Bundles

6.5 Predictive power of the present-value model

The predicted values of the ca variable are calculated from a finite version of equation (6.33), where the number of forecasts ahead is set as the maximum number of periods needed for all the system variables to flatline and all the forecasted variables are first demeaned of their long run mean. Results are plotted in Figures 6.15-21. Tests of the present value model of the current account are evidently rejected given the poor fit of the predictions face to the actual values As pointed out in the previous section, the low R^2 of the Dno equation seems to be the main cause of these results.

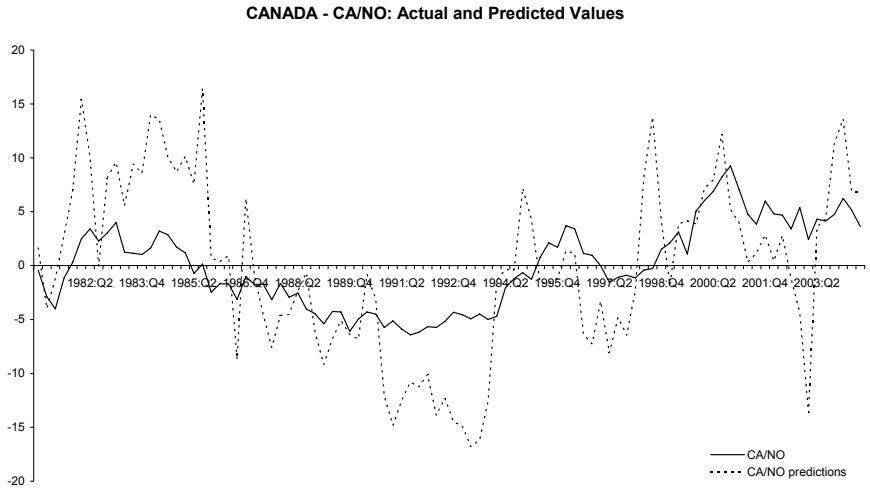


Figure 6.15: Canada - Present Value Model Predictions

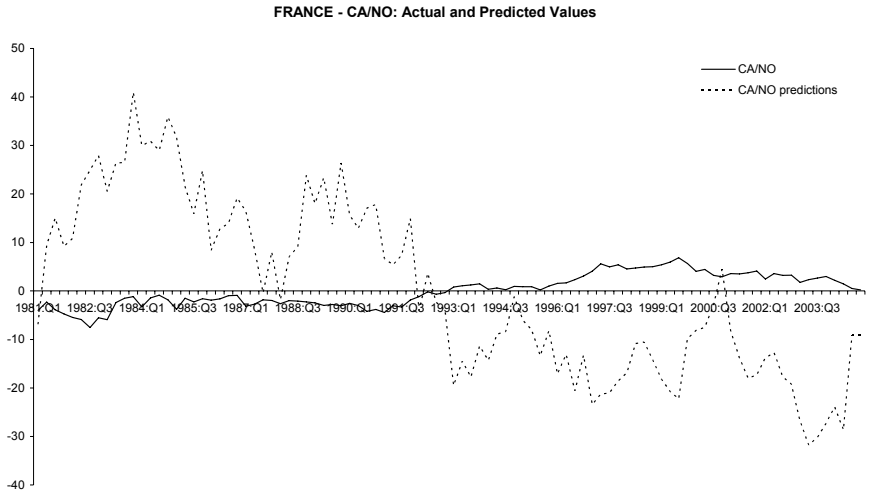


Figure 6.16: France - Present Value Model Predictions

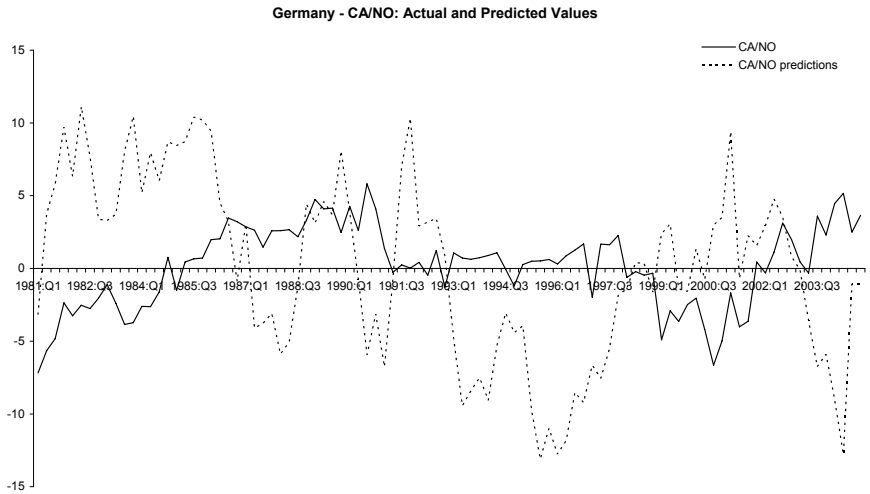


Figure 6.17: Germany - Present Value Model Predictions

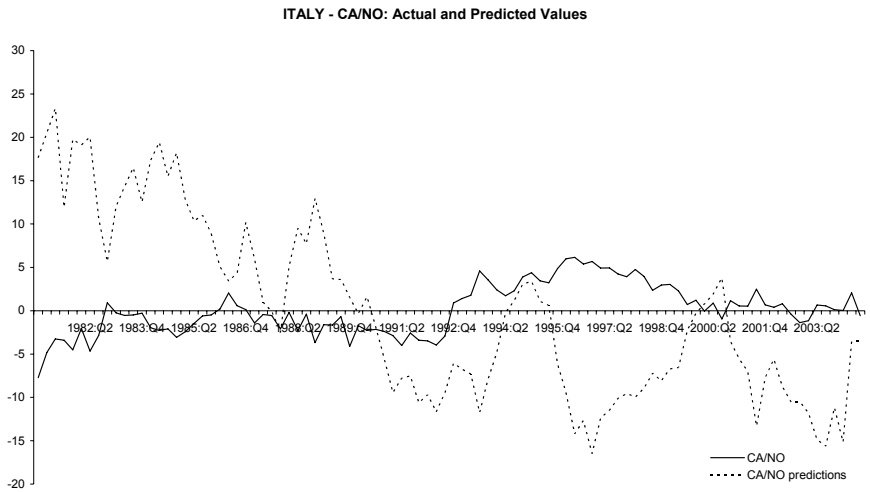


Figure 6.18: Italy - Present Value Model Predictions

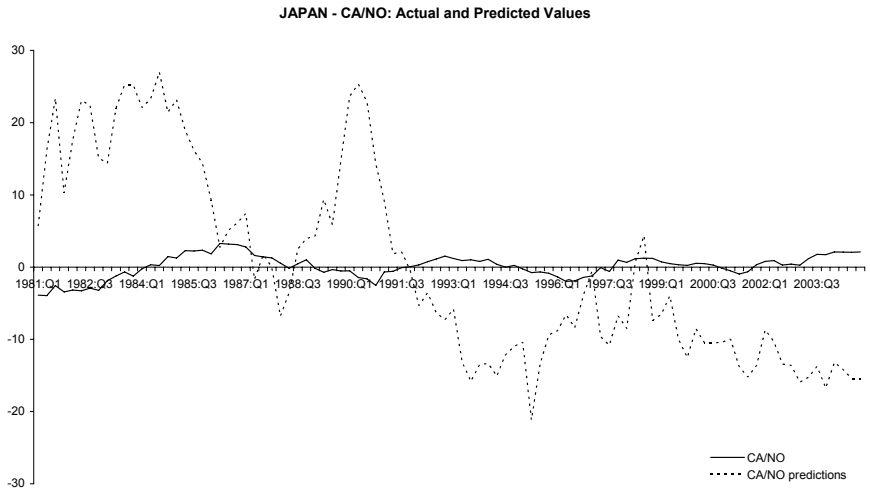


Figure 6.19: Japan - Present Value Model Predictions

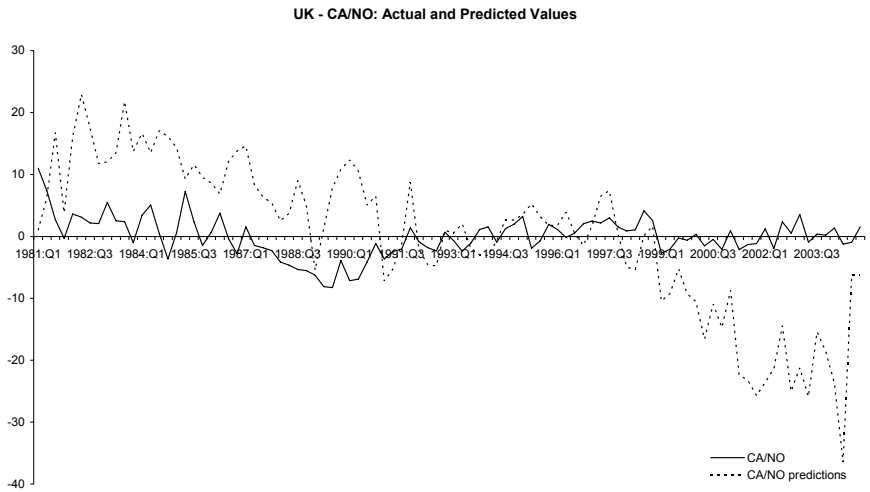


Figure 6.20: United Kingdom - Present Value Model Predictions

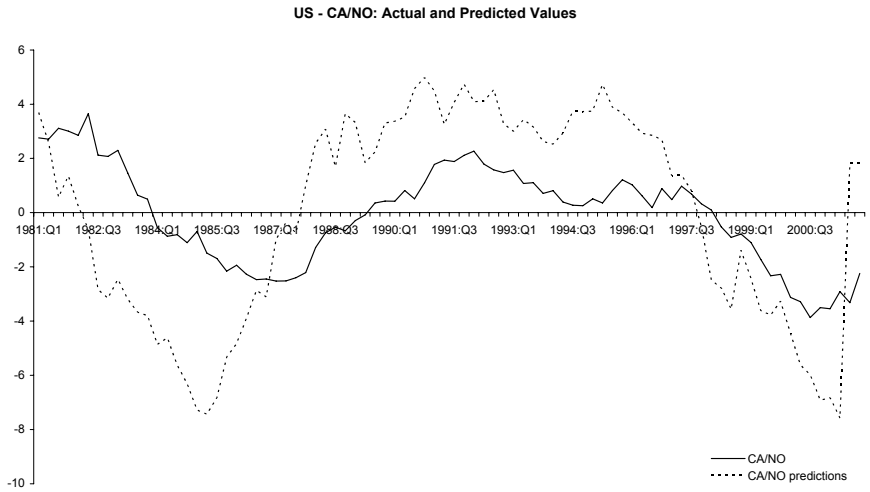


Figure 6.21: United States - Present Value Model Predictions

The present value model of the ca is rejected for all countries also once we compare the variances for the actual and predicted series (reported in Table 6.6). For all G7 countries, the predicted variance is much larger than the actual one, although the existence of a measurement error would predict exactly an opposite result. The extra volatility of

the predicted series might be due to the inclusion of the time varying exchange rate in the model.

Variances

Country:	*CA/NO Actual Data*	*CA/NO Predicted Value*
Canada	15.15	65.83
France	10.78	326.32
Germany	7.62	39.03
Italy	8.23	107.70
Japan	2.49	183.11
UK	10.44	151.29
US	3.34	14.62

Table 6.6: Variance Comparison

Engel and Rogers (2006), with reference to the US economy, also find that the dynamics of its current account are difficult to explain under a particular statistical model of expectations of future US growth. Nevertheless, their results improve once actual data are replaced by survey data on forecasted GDP growth in the G7, as a way of including in the model all the information available to agents and therefore their resulting consumption decisions. However, when we try to improve the fit of the predicted values by using forecasters' estimations for Dno we do not significantly improve our results. We replace our model forecasts for Dno with estimates of future GDP shares based on survey data from Consensus Forecasts, a publication of Consensus Economics Ltd. Data are available on a biannual basis as from 1989Q4. For brevity, we omit to report the estimation results.

6.6 Concluding remarks

In this chapter we run a nonlinear test of the present value model of the current account. We estimated a multivariate system for current account, net output, world real interest rate and real exchange rate, allowing for a threshold adjustment in the current account and exchange rate variables. We then used Monte Carlo integration methods to compute forecasts of these variables over a finite sample, as suggested by the speed of adjustment of their temporary components. Previous evidence on the threshold behaviour of the current account and on the significant effect of the real exchange rate on its adjustment, as presented in the previous chapters, seemed to support this type of methodology.

Nevertheless, in our analysis of the intertemporal model of the current account, the predictive power of our forecasts remain highly unsatisfactory. The difficulty in modelling the net output variable and its resulting poor fit represents the main difficulty. As an attempt to overcome this obstacle, we try to replace our system forecasts by actual forecasters'survey data but we obtain only a marginal improvement in the current account predictions.

6.A Appendix: Data and variable definitions

We look at quarterly data for the period 1979:1 to 2004:4. Following Chinn and Lee (2005), we decide to estimate the US series over a shorter sample period, up to 2001Q4, in order to avoid the substantial change in the country econometric relationship in the post-2001 period. The data in the analysis are obtained from the International Financial Statistics Database by the IMF. All variables are seasonally adjusted and expressed in national currency. According to national account statistics, the current account variable is estimated as the sum of net exports and net primary income from abroad (NPIA); net output is obtained by subtracting Government consumption expenditure and gross fixed capital formation (investment) to GDP. For the exchange rate variables, we refer to trade weighted real effective exchange rate indices, whereas a decrease in the index corresponds to a depreciation (data are available as from 1980:1). The world real interest rate is computed as the average of the G7 short term real interest rates, deflated by forecasts of inflation from an AR(6) process and weighted by the time-varying share of each country GDP in the G7 total. Data on net asset positions, used in the calibration of the b parameter in the PVM equation, are obtained from the International Investment Position series reported in the IFS database.

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