International Macroeconomic Fluctuations, Capital Mobility and the Current Acccount: A Cointegrated Approach

by

Mathias Hoffmann

Submitted to the Department of Economics
in partial fulfillment of the requirements for the degree of

Doctor of Economics

at the

EUROPEAN UNIVERSITY INSTITUTE

June 1999

Signature of Author . . . .
Department of Economics
July 12, 2004

Accepted by . . . .

Contents

1

2

<center>3</center>

0.1 Acknowledgements

It was Axel Weber who initiated me to macroeconomics and applied econometrics. Without him, this thesis would not have been written. As my Master's thesis advisor at the University of Bonn, he first drew my attention to the European University Institute and wrote letters of recommendation. His enthusiasm for economics continues to inspire me and ever since have I benefited from his professional and personal advice. I would like to thank him for everything.

Throughout my time at the Institute, the Department of Economics provided a uniquely stimulating environment for my work. My thesis advisors Michael Artis and Søren Johansen have done a great job in exposing me to new intellectual challenges while at the same time keeping me on track when I was at risk to go astray and cheering me up when I was doubting the value of my work. I am very grateful to both.

The Institute's most important resource are its students and I would like to thank many of my peers for their feedback and help. As it is impossible to list all of them I will not mention any names but those whom I mean will know.

Secretarial support over these four years was simply excellent. Jacqueline's and Jessica's skills often made the impossible possible and both always had an open ear for problems.

A part of this thesis was written while I was visiting the Haas School of Business and the Department of Economics at the University of California at Berkeley. I would like to thank both Institutions for their hospitality. Seminar participants at UC Berkeley and San Diego, the Universities of Bologna, Exeter and Southampton as well as conference participants at the 1999 Trondheim conference on 'Macroeconomic Transmission Mechanisms', the Royal Economic Society Conference 1999 and the 1998 Econometric Society European Meeting have commented on various stages of the work presented here.

Finally and in particular I would like to thank Sascha Becker, Olaf Unteroberdörster, Christian Upper and Steffen Hörnig for their friendship and my parents, Rosalinde and Hans-Werner, for their love and patience.

0.2 Numbering Conventions for Figures and Tables

Figures and Tables are numbered within each chapter and there are no cross-references between chapters. Figures also have a chapter heading, e.g. figure 5 in chapter 2 will be numbered as Figure 2-5 but will be referenced in chapter 2 as 'figure 5' only.

For technical reasons, in chapter 4, table numbers also have a chapter heading and are referenced with this heading.

Chapter 1

Introduction: A Cointegrated Approach to the Current Account

Keep it simple: as simple as possible, but no simpler.

Albert Einstein

1.1   Scope of the thesis

In this thesis cointegrated vectorautoregressions are used to explore the empirics of the intertemporal approach to the current account recently popularized by Sachs (1981), Obstfeld (1986), Obstfeld and Rogoff (1995a,b), Razin (1995), and Obstfeld and Rogoff (1996). The theoretical framework will be given throughout by quadratic models that allow for simple closed-form solutions. These models are not of particular theoretical interest but rather serve to motivate an important reduced-form implication that should also survive in more complicated model settings: the current account should be an order of magnitude less persistent than its driving forces, savings and investment. This prediction can be formalized as a cointegrating restriction in VAR-approximations of the data dynamics.

Cointegrated models give rise to natural classifications of variables into permanent

6

and transitory components. All chapters exploit this feature as a convenient identification device. In particular, Chapter 3 exploits the error-correction behaviour of such systems to study international capital mobility and the Feldstein-Horioka (1980) puzzle.

The intertemporal theory of the current account also makes strong predictions about the role of country-specific and global shocks. In particular it emphasizes that the current account serves as a buffer that allows the smoothing of consumption through international borrowing and lending. However, in response to global shocks, there is no scope for international borrowing and lending and therefore current accounts should not react to them. In two of the papers, we exploit this prediction of the theory to identify global and country-specific shocks. Chapter 4 studies how country-specific and global shocks identified in this way map into permanent and transitory disturbances identified through the cointegration properties of the reduced form and uses this mapping to derive implications for current account dynamics.

Chapter 2 prepares the scene in that it exploits the special character of the shocks we are out to identify: It assesses the quality of the identification of country-specific and global shocks using cross-country evidence and examines the role of the current account in trend output growth.

Even though each individual paper addresses an economic issue and tries to add to economic knowledge, all three are of methodological interest in that they try to illustrate the role of the appropriate amount of economic theory in empirical economic research. I define a theory as the core of assumptions and predictions that is common to a certain class of models. Each model formalizes the theory but each does so in a different way, using different auxiliary assumptions or refinements. The argument will be that an individual theoretical model should not be taken too seriously. Rather reduced-form implications, common to all or at least most models formalizing the theory should be emphasized. Analysis of the reduced form should then proceed focussing on these key implications, allowing the econometric model to be specified in statistical accordance with the data dynamics. Our claim is that this approach avoids 'measurement without

7

theory' but also counters the problems of overparameterized models that can give rise to empirical puzzles.

We first present the intertemporal approach to the current account in the context of recent developments in dynamic macroeconomics. Another section relates to the methodological insights we hope to illustrate in the chapters of this thesis. As a conclusion to this introductory chapter we provide a synopsis of the thesis by means of chapter abstracts.

1.2   The intertemporal approach

The intertemporal approach is based on the presumption that the properties of macroeconomic aggregates can be approximated by the behaviour of a representative agent that maximizes a discounted stream of period utilities generally derived from consumption. The agent's expectations are consistent with the economic model in that they coincide with the conditional expectations of variables given the model's structure. This assumption is generally referred to as the 'rational' expectations hypothesis (Muth (1961)), even though it is not an assumption about economic rationality but rather about the internal consistency of the model (see the relevant chapters in Hendry (1995)).

Initiated by Lucas (1973), Lucas and Rapping (1969) and promoted by Kydland and Prescott (1977) and others, the rational expectations revolution soon triggered the development of stochastic dynamic general equilibrium models. By emphasizing the intertemporal consistency of economic decision-making, these models created an important rival to the traditional IS-LM paradigm and finally largely replaced it as a framework for macroeconomic analysis.

International macroeconomics long stood apart from this development, continuing to make extensive use of the traditional toolkit. As Krugman (1995) points out, there is a variety of reasons for this: whereas closed-economy macroeconomics can rely on a relatively large body of stylized facts, empirical regularities concerning the cross-country behavior of economic aggregates are much harder to establish and to theorize on. Or, if

8

they exist, they are difficult to reconcile with the type of intertemporal models usually employed in closed-economy macroeconomics. This gives rise to numerous puzzles, of which the consumption correlation puzzle, the Feldstein-Horioka puzzle, the home bias puzzle or the real exchange rate puzzle are just some of the most prominent examples.

The existence of numerous puzzles can also be interpreted as a corollary of the intellectual state of international macroeconomics described by Krugman as one of 'intellectual distress': puzzles can arise if the body of theory employed is very heterogenous and conflicting. Krugman highlights three missing theoretical links in international macroeconomics:

1. The lack of trade-theoretical foundations in models of international finance: theories about a country's long-term external adjustment are generally not consistent with workable models that allow one to address issues of short-run exchange-rate and balance-of-payments adjustment.

2. Whereas the simplistic framework of IS-LM/ Mundell-Fleming seems to work remarkably well in guiding policy decisions, intertemporal models generally fail to produce robust policy advice and have high informational requirements about the structure of shocks.

3. The difficulty in reconciling rational expectations with the observed stickiness of nominal variables, whereas in international macro there is overwhelming evidence that prices are sticky (see e.g. the literature on the real exchange rate puzzle surveyed in Rogoff (1996)).

However, a coherent framework for most major issues in international finance and macroeconomics has started to emerge over the last decade. Like closed-economy macroeconomics it is based on intertemporal optimisation by a representative economic agent and it emphasizes the role of the current account as the main variable in the international proliferation of economic impulses. Early contributions to this literature go back to Sachs (1981) and Obstfeld (1986). The *intertemporal approach to the current account*

9

has become a standard paradigm in recent years when it became apparent that it could be reconciled with sticky-price features so important in international macro (Obstfeld and Rogoff (1995a)) and with models of international trade in the spirit of Dornbusch, Fischer and Samuelson (1977). Indeed, the book by Obstfeld and Rogoff (1996) demonstrated for the first time how a wide range of apparently disparate topics in international macro and finance and trade could all be addressed within a coherent model framework, essentially resolving the first and third of Krugman's linkage problems. Much of the motivation for this thesis is taken from the Obstfeld-Rogoff book.

Whereas many theoretical issues have been resolved, the empirics of the intertemporal approach to the current account has attracted much less attention. This is true in particular for formal econometric testing, even though there are a few notable exceptions including the papers by Sheffrin and Woo (1990) and Gosh (1995) who test the present-value theory of the current account implied by the intertemporal approach.

Only very few contributions to the literature have moved on to investigate further implications of the theory, in particular its strong predictions about the role of country-specific and global shocks and about the role of different degrees of persistence for the dynamics of the current account. Glick and Rogoff (1995) estimated a structural econometric model derived from the explicit linearization of an intertemporal model. They find that mainly country-specific shocks drive the current account - in accordance with the theory. Rogers and Nason (1998) use a structural VAR approach and employ various identification schemes. They find their results to be highly sensitive to perturbations in the identification scheme. In particular, Choleski-type identifications are found to yield long-run dynamics that are inconsistent with long-run identification schemes in the spirit of Blanchard and Quah (1989) and vice versa.

10

1.3 Some methodology

Modern macroeconomic theorizing is largely motivated by the effort to rationalize so called 'stylized facts', i.e. statistical properties of the data that are found to be robust over time and across different data sets. On the other hand, macroeconomic theory itself, starting from *a priori* specifications, can sometimes generate strong predictions for the stylized behaviour of economic data. That stylized facts - or more generally: statistically robust features of the data - should be rationalized by economic theories and that in turn theories should be tested against the data- this is probably the way most economists could agree upon in which macroeconomics as a science should proceed.

However, there is a lot of disagreement in the profession about the exact way and the role of *a priori* economic theory in macroeconometrics. The volume edited by Hoover (1995), contrasts the tensions of the field by juxtaposing contributions from some of the most prominent proponents of the different school of thoughts. To bring forward my argument, I will focus on what I think are the two most important approaches - the theory-driven and the data-driven one.

I mean by the 'theory-driven' or 'American' approach to macroeconomics that macroeconomic model-building starts from trying to emulate stylized facts through model calibration. This largely coincides with the RBC-literature exemplified in the work of Kydland and Prescott (1982), Long and Plosser (1983) and forcefully advocated in Prescott (1986). The reduced form of the model is derived from first principles including optimising behaviour, rational expectations and market-clearing. The parameters of the model are then chosen in such a way that model simulation will on average mimic some moments of interest of actual economic aggregates. Whereas this procedure can be useful as an exercise in quantitative economic theorizing, calibration does not amount to rigorous econometric testing as only a subset of all data moments can be matched to the data. Also, econometric identification of the structural parameters will generally not be possible or requires imposing untestable just-identifying restrictions - an immediate consequence of the overparameterization of this type of models.

11

On the other hand, empirical macroeconomics over the last two decades has adopted more and more sophisticated time series methods. In particular the development of the statistical theory of non-stationary processes has been motivated by the recognition that many macroeconomic time series display substantial persistence. Macroeconomic aggregates are described by stochastic processes, in terms of their degree of integration, stationary (cointegrating) relations that prevail between them and the like. The model is required to use all statistical information, whereas *a priori* economic theory is employed mainly at a low level to determine the choice of information set. I will refer to this approach as the 'data-driven' or 'European' approach to macroeconomics. It is exemplified in the work of Johansen and Juselius (1990), Hendry and Mizon (1990) and others and forcefully restated in Hendry (1995).

Unfortunately, the 'data-driven' and the 'theory-driven' approaches do seem orthogonal to each other in terms of their respective languages and concepts and results found in one are not easily translated into the other approach. Levtchenkova and Pagan (1998) and Juselius (1999) argue, that even to the degree that the two approaches share a common language, similar sounding jargons are treacherous in that they conceal profound differences.

In this thesis we wish to illustrate the following claims:

- The non-stationary character of many macroeconomic aggregates should not be regarded as a nuisance but rather as a useful identifying device. The same economic theory when expressed under the assumption that the data to be modelled are non-stationary can create much stronger empirical predictions than when no assumptions about persistence are made.

  This is the common insight that underlies some recent important breakthroughs in macroeconomic modelling.

  1. The King-Plosser-Stock-Watson (1991) approach: in its simplest version, the insight of the King-Plosser-Stock Watson approach is that a basic stochastic

12

growth model predicts that output, investment and consumption have a common trend. Hence, there are two cointegrating relationships between them. These are the great ratios: investment over output and consumption over output should be fairly constant over time. King's et al. argument is that the cointegration relations predicted by the theory are not only important for the test of balanced growth theory but also for business cycle modelling: Valid inference about the higher-frequency dynamics is possible only once the long-run structure is adequately modelled. But inference about the long-run structure requires the non-stationarity of the data.

2. Present-value models and cointegration (Campbell and Shiller (1987)):

Rational expectations models quite often give rise to a present-value formula which relates the spread (i.e. the difference) between to macroeconomic aggregates to the discounted sum of expected future changes of a driving variable.

Whereas such a present-value relation is usually not easy to test formally, it has a straightforward implication once the two macroeconomic aggregates are characterized as integrated processes. Then the present-value formula predicts that the spread is the discounted sum of a differenced integrated process. In other words, the spread is stationary whereas the two macroeconomic aggregates individually are not. Hence, there is a cointegrating relation between the two aggregates.

This finding by Campbell and Shiller has made an enormous impact on empirical modelling in finance and macroeconomics: the term-structure of interest rates, option and stock pricing, the macroeconomic dynamics of consumption and the current account as well as the analysis of fiscal solvency - all these issues have afterwards been addressed in a cointegrated framework.

- Some apparent puzzles that arise in the framework of the 'theory-driven' approach can be better understood or even resolved once the non-stationary character of

13

macroeconomic data is recognized. The economic theory should actually be understood as a theory that is conditional on past information. It is a crucial step to recognize that the data forming the information set is non-stationary and that this non-stationarity has to be accounted for in formulating the reduced-form implications of the economic theory. In particular, this implies that empirical modelling has to allow for adjustment lags, error correction and the like. Unless the dynamics of the data given the reduced-form model are appropriately specified, there is little hope that the model can be used for theory check and theory development.

- To the degree that integratedness of the data is a convenient identifying device rather than a nuisance - as argued above - we should not understand persistence in an absolute way. The unit root we find in a macroeconomic time series is not the truth. It is rather a convenient classification of this time series as 'relatively persistent', where 'relative' pertains to the information set that economic theory tells us is relevant. This way of reasoning has two implications: first, we should think hard about which time horizon the economic theory we are out to test actually is meant to apply. Depending on the time horizon, it may prove useful to treat an economic variable as persistent in one theoretical context but as stationary in another.

  Secondly, it also means that univariate time-series properties, in particular unit-root tests, are not particularly meaningful. An economic theory usually is about several variables and a classification into 'stationary' and 'persistent' should be undertaken vis-a-vis this information set.

We will now provide a short overview of the three essays that form this thesis. Whereas all three illustrate the points aforementioned, the reader will realize that each in itself is not meant primarily as an illustration of econometric methodology but is driven by an economic problem.

Whereas the theoretical framework I use is an 'American' one in the sense of the

14

definitions given above, the econometric methods I employ are typically associated with the 'European' approach. Even though the language of the two approaches sounds quite often very similar, the logic is often different to a degree that no one-to-one analytical mapping between a theoretical model and the reduced form can be derived. It is there where economic judgement has to come in and where I have to recur to analogies: does a feature of the theoretical model have any correspondence in the reduced-form? If so, can we test for this corresponding feature or does it have any meaning at all in the reduced-form setup?

This translation exercise requires using more *a priori* theory than is typically employed in the 'data-driven' approach but less theory than is usually employed in the 'theory'-driven approach. The methodological stance I take comes close to the one taken by Canova (1995) in the volume edited by Hoover (1995):

> 'A VAR econometrician can be thought of as a rational expectations econometrician who is skeptical of many of the restrictions that a particular formulation of dynamic economic theory imposes [...]. Therefore, in order to produce a structural interpretation of the VAR model, he uses only a limited number of these constraints and "lets the data speak" [...].' p. 68

Against this background, the chapters of this thesis can also be read as an effort to bridge or at least highlight the language differences between the two approaches and to illustrate some implications for theorizing and measurement. I do not consider this effort exhaustive nor do I make any claim as to its final success.

1.4   Chapter Abstracts

Here we provide abstracts for the three papers of this thesis.

1.4.1 National Stochastic Trends and International Macroeconomic Fluctuations: the Role of the Current Account

We propose a simple intertemporal model of output and current account dynamics that we estimate using a cointegrated VAR approach. We suggest a method for identifying global and country-specific shocks from the VAR and test it, using cross-country evidence. Our results show that the identification scheme works well in practice, corroborating an important prediction of the intertemporal approach to the current account. We associate global shocks with movements in the US output growth rate. In accordance with the theory, we also observe a link between the global shock and a measure of the world real interest rate. This link is more pronounced in the long-run than in the short-run.

1.4.2 The Feldstein-Horioka Puzzle and a New Measure of International Capital Mobility

In intertemporal optimization models of current account dynamics, the budget constraint will induce high degrees of positive comovement in the levels of savings and investment and the two variables are likely to be cointegrated. Error correction will then also influence the correlations of the cyclical components which are per se uninformative about capital mobility. As an alternative we suggest a new measure of long-run capital mobility based on Johansen's procedure. We apply our method to historical British and US data and find surprisingly high levels of long-run capital mobility throughout the century.

1.4.3 Current Accounts and the Persistence of Global and Country-Specific Shocks: Is Investment really too Volatile?

Using a small VAR of the current account and investment, we identify two categories of shocks: permanent vs. transitory and country-specific vs. global. Our approach involves only the most minimal identifying assumptions. Using data from the G7 countries, we

16

find that some important predictions of the intertemporal approach to the current account are confirmed by the data. We are also able to solve the puzzle encountered by Glick and Rogoff (1995) that the investment response to country-specific shocks is excessive vis-a-vis the current account response: the estimated response is an amalgam of responses to permanent and transitory shocks. In our specification the current account reacts as predicted to the permanent component of country-specific shocks and we find investment not to be excessively volatile.

17

Bibliography

[1] Blanchard, Olivier and Quah, Danny, (1989), "The Dynamic Effects of Aggregate Demand and Supply Disturbances," *American Economic Review* 79: 655-73.

[2] Canova, F. (1995), 'The Economics of VAR models', in Hoover, K. (ed.) *Macroeconometrics - Developments, Tensions and Prospects*, Kluwer Academic Publishers.

[3] Campbell, J. Y, and Shiller, R.J., (1987), ' Cointegration and tests of present value models,' *Journal of Political Economy* 95: 1062-1088.

[4] Dornbusch, R. Fischer, S. and Samuelson, P. A. (1977), 'Comparative advantage, trade and payments in a Ricardian model with a continuum of goods' *American Economic Review*, 67: 823-39.

[5] Ghosh, A., (1995), 'Capital mobility amongst the major industrialized countries: Too little or too much?,' *Economic Journal* 105: 107-128

[6] Glick, R. and Rogoff, K., (1995), 'Global versus country-specific productivity shocks and the current account,' *Journal of Monetary Economics,* 35, pp. 159-92.

[7] Hendry, D. (1995), *Dynamic Econometrics*, Oxford University Press.

[8] Hendry, D. and Mizon, G. (1990), 'Evaluating Dynamic Econometric Models by Encompassing the VAR.' in P. C. B. Phillips and V.B. Hall (eds.) *Models, Methods and Applications of Econometrics: Essays in Honour of Rex Bergstrom* Oxford: Oxford University Press.

[9] Hoover, K. (ed.) (1995) *Macroeconometrics - Developments, Tensions and Prospects*, Kluwer Academic Publishers.

[10] Johansen, S. and Juselius, K.(1990), 'Maximum likelihood estimation and inference on cointegration - with application to the demand of money', *Oxford Bulletin of Economics and Statistics*, 52, pp. 169-210.

[11] Juselius, K. (1999), 'Models and Relations in Economics and Econometrics', *Journal of Economic Methodology*, forthcoming.

[12] Krugman, P. (1995), 'What do we need to know about the international monetary system?' in Kenen, P. (ed.) *Understanding Interdependence*, Princeton University Press.

[13] King, Robert, Plosser, Charles I., Stock, James H. and Watson, Mark W., (1991), "Stochastic Trends and Economic Fluctuations", *American Economic Review* 81: 819-40.

[14] Kydland, F. and Prescott, E. (1977), 'Rules rather than discretion: the inconsistency of optimal plans', *Journal of Political Economy* 85, 3 (June), pp. 473-92.

[15] Kydland, F. and Prescott, E. (1982), 'Time to build and aggregate fluctuations', *Econometrica* 59: 1345-70.

[16] Levtchenkova, S. and Pagan, A. (1998), 'Shocking Stories', *Journal of Economic Surveys* 12:507-32.

[17] Long, J. B. and Plosser, C. I. (1983), 'Real business cycles', *Journal of Political Economy* 91: 39-69.

[18] Lucas, Robert E. (1973), 'Some International Evidence on Output-Inflation Trade-offs', *American Economic Review* 63: 326-34.

[19] Lucas, R. E. and Rapping, L. (1969), 'Real wages, employment and inflation', *Journal of Political Economy* 77: 721-54.

[20] Muth, John F. (1961), 'Rational Expectations and the Theory of Price Movements', *Econometrica* 29: 315-35.

[21] Nason, J. and Rogers, J. H. (1998), 'Investment and the Current-Account in the Short-Run and the Long Run', mimeo, University of British Columbia.

[22] Obstfeld, M., (1986), 'Capital Mobility in the World Economy: theory and measurement', *Carnegie Rochester Conference Series on Public Policy* 24: 55-104.

[23] Obstfeld, M. and Rogoff, K. (1995a), 'Exchange Rate Dynamics Redux', *Journal of Political Economy* 103: 624-60.

[24] Obstfeld, M. and Rogoff, K. (1995b), 'The Intertemporal Approach to the Current Account', in Gene M. Grossman and K. Rogoff, (eds.), *Handbook of International Economics*, vol. 3. Amsterdam: North Holland.

[25] Obstfeld, M. and Rogoff, K., (1996), *Foundations of International Macroeconomics*, MIT Press, Cambridge.

[26] Prescott, E. C. (1986), 'Theory ahead of business-cycle measurement' *Carnegie-Rochester Conference Series on Public Policy* 25: 11-44.

[27] Razin, A. (1995), 'The Dynamic-Optimizing Approach to the Current Account: Theory and Evidence', in Kenen, P. (ed.) *Understanding Interdependence,* Princeton.

[28] Rogoff, K. (1996), 'The purchasing power parity puzzle', *Journal of Economic Literature* 34: 647-88.

[29] Sachs, J, (1981), 'The Current Account and Macroeconomics adjustment in the 1970s,' *Brookings Papers on Economic Activity* 1:201-268.

20

[30] Sheffrin, S. and Woo, W. T. (1990), 'Present Value Tests of an Intertemporal Model of the Current Account' *Journal of International Economics* 29: 237-53.

[31] Summers, L. H. (1991), 'The Scientific Illusion in Empirical Macroeconomics', *Scandinavian Journal of Economics* 93: 129-48.

21

Chapter 2

National Stochastic Trends and International Macroeconomic Fluctuations: the Role of the Current Account

2.1 Introduction

Little stylized knowledge is available on the question in which way industrialized countries are prone to international shocks and how they adjust to them. In this paper, we propose a simple model centered around the current account as the key variable of macroeconomic transmission. Our setup offers a compact framework in which the following questions can be tackled:

- Can we validly identify global and country-specific shocks using a simple model of the world economy?

- How persistent are global and country-specific shocks?

- Can we associate global shocks with observable economic variables?

22

- What drives the development of long-run output in the seven biggest economies? Is it global shocks or country-specific shocks? Do shocks to the current account drive output or do output shocks determine the current account?

The theoretical framework of the paper is provided by the intertemporal approach to the current account initiated by Sachs (1981) and extended by Obstfeld (1986, 1995). Since the appearance of the landmark book by Obstfeld and Rogoff (1996), the intertemporal approach has also become a textbook paradigm. Our empirical implementation relies on a structural VAR approach that is embedded in a cointegrated model. We think that such a framework is a good vehicle with which to fish for stylized facts in international macro: it contains enough economics to avoid the risk of 'measurement without theory' but is at the same time simple and data-driven.

The paper's layout is as follows: section two presents a simple intertemporal optimisation model of the current account that highlights the econometric implications of the intertemporal approach and suggests how permanent and transitory components of output can be identified. In Section 3, we suggest an identification scheme to identify country-specific and global shocks and discuss its econometric implementation. In Section 4, we present results; in particular, we discuss the quality of our identification scheme, using cross-country evidence. Section 5 concludes.

2.2   The intertemporal approach

In our empirical implementation, we will use expected utility, which is quadratic in consumption, in an intertemporal setting: i.e. the representative consumer maximizes

$$E_t \sum_{i=0}^{\infty} \left(\frac{1}{1+r} \right)^i \left[C_{t+i} - \frac{h}{2} C_{t+i}^2 \right] \tag{2.1}$$

23

subject to the intertemporal budget constraint

$$B_{t+1} = (1+r)B_t + Y_t - C_t \tag{2.2}$$

where Y_t is output, C_t is consumption and r represents the world real interest rate. B_t denotes the stock of net foreign assets which is required to be non-explosive:

$$\lim_{i \to \infty} B_{t+i}(1+r)^{-i} = 0 \tag{2.3}$$

The current account is defined as[1]

$$CA_t = \Delta B_{t+1} \tag{2.4}$$

In such a model agents behave as if all variables actually realize their expected values. This certainty-equivalence feature yields a simple forward looking solution for the consumption function:

$$C_t = \frac{r}{1+r}\left[(1+r)B_t + \sum_{s=0}^{\infty}\left(\frac{1}{1+r}\right)^s E_t Y_{t+s}\right]$$

Plugging this into the definition of the current account, we get

$$CA_t = Y_t - \frac{r}{1+r}\sum_{s=0}^{\infty}\left(\frac{1}{1+r}\right)^s E_t Y_{t+s} = Y_t - \tilde{Y}_t \tag{2.5}$$

where $\tilde{Y}_t$ denotes the permanent value of output.

Now let us specify a simple process for output:

---

[1]In this model, a change in the net foreign asset position, B_t, will require an international flow of funds. The current account is more generally defined as the difference between savings and investment, $CA = S - I$ and of course that is the case here as well once we define $S_t = Y_t - C_t + rB_t$. The equality between CA_t and ΔB_{t+1}, will hold only under the assumption that no price changes affect the country's net foreign asset position. This would, e.g., happen whenever the real exchange rate changes.

24

$$Y_t = Y_{t-1} + \sum_{i=0}^{\infty} \mathbf{c}_i' \mathbf{e}_{t-i} \tag{2.6}$$

Here, $\mathbf{e}_t = \begin{bmatrix} e_t^c, & e_t^w \end{bmatrix}'$ denotes the vector of country-specific and global shocks which are assumed to have unit variance and are serially and contemporaneously uncorrelated.

We can rewrite equation (2.5) to yield:

$$CA_t = -\sum_{s=1}^{\infty} \left(\frac{1}{1+r}\right)^s E_t \Delta Y_{t+s} \tag{2.7}$$

Then, from (2.6) we get

$$E_t \Delta Y_{t+s} = \sum_{i=0}^{\infty} \mathbf{c}_{i+s}' \mathbf{e}_{t-i}$$

Plugging this into (2.7) yields:

$$CA_t = -\sum_{s=1}^{\infty} \left(\frac{1}{1+r}\right)^s \sum_{i=0}^{\infty} \mathbf{c}_{i+s}' \mathbf{e}_{t-i} = -\sum_{i=0}^{\infty} \mathbf{d}_i' \mathbf{e}_{t-i}$$

where $\mathbf{d}_i' = \sum_{s=1}^{\infty} \left(\frac{1}{1+r}\right)^s \mathbf{c}_{i+s}'$.

The above setup gives us a simple joint representation of current account and output in differences:

$$\begin{bmatrix} \Delta CA_t \\ \Delta Y_t \end{bmatrix} = \begin{bmatrix} (\mathbf{1} - \mathbf{L})\mathbf{d}'(\mathbf{L}) \\ \mathbf{c}'(\mathbf{L}) \end{bmatrix} \mathbf{e}_t = \mathbf{D}(\mathbf{L})\mathbf{e}_t \tag{2.8}$$

Note that in this structural moving-average representation, the dynamics of the current account are driven by global and country-specific shocks. If however, international capital mobility is sufficiently high, all countries will react to a global shock in the same way - wanting to save more or less, depending on which way the shock goes. But not all can

25

Hoffmann, Mathias (1999), International Macroeconomic Fluctuations, Capital Mobility and the Current Account: A cointegrated approach
European University Institute
DOI: 10.2870/45469

run current account deficits or surpluses at the same time. Rather, a global shock should then impinge on the world interest rate and equilibrate world saving and investment.

This reasoning has two implications:

- The current account should react more strongly to country-specific shocks than to global shocks.

- global shocks should be associated with changes in the world interest rates.

In the sequel of the paper, we will use the first of these two implications to identify country-specific and global shocks. The quality of this identification is then assessed using the second.

2.3   Econometric Implementation

In the structural model (2.8), both variables are stationary. In this paper, however, we are concerned with the long-run properties of output, i.e. with its permanent component. We will therefore consider a system in the level of output and the current account:

$$X_t' = \left[\begin{array}{cc} CA, & Y_t \end{array} \right] \tag{2.9}$$

In such a system, output is $I(1)$ whereas the current account is stationary. This amounts to saying that the two variables share one common trend or in other words, there is a trivial cointegrating relationship with cointegrating vector $\beta' = \left[\begin{array}{cc} 1, & 0 \end{array} \right]$. This, becomes clearer once we express X_t in terms of a (structural) Beveridge-Nelson (1981)/Stock-Watson (1988) representation:

$$\mathbf{X}_t = \mathbf{D(1)} \sum_{i=0}^{t} \mathbf{e}_i + \mathbf{D^*(L)e}_t \tag{2.10}$$

where $\mathbf{D}_i^* = -\sum_{l=i+1}^{\infty} \mathbf{D}_l$ and $\mathbf{D(1)} = \sum_{i=1}^{\infty} \mathbf{D}_i$.

26

Because CA is stationary, we have $\mathbf{d}'(\mathbf{1}) = 0$ and therefore

$$\mathbf{D}(\mathbf{1}) = \left[\begin{array}{c} \mathbf{d}'(\mathbf{1}) \\ \mathbf{c}'(\mathbf{1}) \end{array} \right] = \left[\begin{array}{cc} 0 & 0 \\ c_{CA}(1) & c_Y(1) \end{array} \right]$$

Hence, $\mathbf{D}(\mathbf{1})$ has reduced rank and the long-run dynamics of the system are driven by the stochastic trend $\mathbf{c}'(\mathbf{1}) \sum_{i=0}^{t} \mathbf{e}_i$.

The structural shocks are unobservable and therefore the moving average-representation of $\Delta\mathbf{X}_t$ or the BN-representation for $\mathbf{X}_t$ cannot be estimated directly. Rather, we assume that it is possible to estimate a reduced-form moving average

$$\Delta\mathbf{X}_t = \mathbf{C}(\mathbf{L})\varepsilon_t \tag{2.11}$$

In which the only way the global and country-specific shocks get 'mixed up' is that they are a linear combination of the reduced-form residuals:

$$\varepsilon_t = \mathbf{S}\mathbf{e}_t \tag{2.12}$$

As we assumed the global and country-specific shocks to be *i.i.d.* and to have unit-variance as well as to be contemporaneously uncorrelated, the variance-covariance matrix Ω of the reduced-form residuals is given by

$$\Omega = \mathbf{S}\mathbf{S}' \tag{2.13}$$

In our two-dimensional system, this condition imposes three restrictions on $\mathbf{S}$. To just identify $\mathbf{S}$, one further restriction is needed.

Theory predicts that the current account should react only weakly to global shocks. We will exploit this property here to disentangle global from country-specific shocks. In so doing, we will impose the restriction that global shocks do not have an effect on

the current account in the period they occur (they can however have a non-zero effect later). In fact, imposing this restriction amounts to a very simple identifying restriction: identification is achieved by means of a Choleski decomposition of the variance-covariance matrix of the reduced form residuals, Ω. To see this, note that the first component of $\boldsymbol{\varepsilon}_t$ is the reduced-form innovation to the current account. Requiring that only country-specific shocks drive this component, we get

$$\mathbf{S} = \left[\begin{array}{cc} s_{11} & 0 \\ s_{21} & s_{22} \end{array} \right] \tag{2.14}$$

But together with $\Omega = \mathbf{SS}'$ this uniquely identifies $\mathbf{S}$ as the lower Choleski-factor of Ω.

Hence, we can map the structural MA-form into the reduced form:

$$\mathbf{C(L)S} = \mathbf{D(L)} \tag{2.15}$$

And as our interest will be particularly in long-run forces:

$$\mathbf{C(1)S} = \mathbf{D(1)}$$

We will now approximate $\mathbf{C(L)}$ by a VAR-representation. Note, however, that a finite-order VAR representation for $\Delta\mathbf{X}_t$ does not exist due to the presence of a common trend. It follows from Granger's representation theorem (Engle and Granger (1987)) that $\Delta\mathbf{X}_t$ can be represented in the form of a vector-error correction model (VECM):

$$\Gamma(\mathbf{L})\Delta\mathbf{X}_t = \boldsymbol{\alpha}CA_{t-1} + \boldsymbol{\varepsilon}_t \tag{2.16}$$

where $\Gamma(\mathbf{L})$ is a 2×2 matrix-polynomial and $\boldsymbol{\alpha}' = \left[\begin{array}{cc} \alpha_1 & \alpha_2 \end{array} \right]$.

Once we have estimated this model, we can express the long-run structure of output as a function of the parameters of the VECM. In particular, as demonstrated in

28

Johansen (1995), the matrix $\mathbf{C}(\mathbf{1})$ can be given a closed-form representation in terms of the parameters of the cointegrated VAR:

$$\mathbf{C}(\mathbf{1}) = \boldsymbol{\beta}_\perp (\boldsymbol{\alpha}'_\perp \Gamma(\mathbf{1}) \boldsymbol{\beta}_\perp)^{-1} \boldsymbol{\alpha}'_\perp$$

Now note that the structure of this matrix is such that it maps the reduced-form disturbances $\boldsymbol{\varepsilon}_t$ into the span of $\boldsymbol{\alpha}_\perp$. The disturbances $\boldsymbol{\alpha}'_\perp \boldsymbol{\varepsilon}_t$ accumulate to the permanent component of $\mathbf{X}_t$ whereas transitory disturbances will be in the null space of $\mathbf{C}(\mathbf{1})$ We can therefore define the permanent disturbances as

$$\eta_t = \boldsymbol{\alpha}'_\perp \boldsymbol{\varepsilon}_t \tag{2.17}$$

and by requiring that permanent and transitory disturbances be orthogonal to each other, we get the transitory shocks as

$$\tau_t = \boldsymbol{\alpha}' \Omega^{-1} \boldsymbol{\varepsilon}_t \tag{2.18}$$

Denoting

$$\boldsymbol{\theta}'_t = \begin{bmatrix} \eta_t, & \tau_t \end{bmatrix} \tag{2.19}$$

we then have $var(\theta) = diag\{var(\eta), var(\tau)\} = \begin{bmatrix} \boldsymbol{\alpha}'_\perp \Omega \boldsymbol{\alpha}_\perp & 0 \\ 0 & \boldsymbol{\alpha}' \Omega^{-1} \boldsymbol{\alpha} \end{bmatrix}$.

In the present bi-variate case with $\boldsymbol{\beta}' = \begin{bmatrix} 1, & 0 \end{bmatrix}$, we have $\boldsymbol{\beta}'_\perp = \begin{bmatrix} 0, & 1 \end{bmatrix}$. Furthermore, $\boldsymbol{\alpha}'_\perp = \begin{bmatrix} -\alpha_2, & \alpha_1 \end{bmatrix}$. Let also $\Gamma(\mathbf{1}) = \{\gamma_{ij}\}_{i,j=1,2}$. Then it is easily verified that $\mathbf{C}(\mathbf{1})$ is of the form

$$\mathbf{C}(\mathbf{1}) = \begin{bmatrix} 0 & 0 \\ c_{21}(1) & c_{22}(1) \end{bmatrix} \tag{2.20}$$

29

where

$$c_{21}(1) = \frac{-\alpha_2}{-\alpha_2\gamma_{12} + \alpha_1\gamma_{22}} \text{ and } c_{22}(1) = \frac{\alpha_1}{-\alpha_2\gamma_{12} + \alpha_1\gamma_{22}} \qquad (2.21)$$

2.3.1   The long-run effects of shocks

In a seminal paper, Blanchard and Quah (1989) identified demand and supply distur-
bances from a bivariate system, requiring that the former do not have a long-run effect on
output. Their restriction postulates a form of long-run neutrality that - in various settings
- is often suggested by economic theory. This is why the Blanchard-Quah identification
scheme has proven very popular in applied work over the last decade (for applications
of the Blanchard-Quah scheme see e.g. Bayoumi and Eichengreen (1992 a and b) and
Bayoumi and Taylor (1995)).

Also in the context of this paper, the Blanchard-Quah identification seems an obvious
candidate. Economic models will often require that country-specific shocks are long-run
neutral w.r.t. output. For example in the Glick and Rogoff (1995) model, the empirical
implementation will yield results that are at odds with the short-run dynamics of the
intertemporal theory if in the theoretical model country-specific total factor productivity
is required to follow a random walk.

In a recent study, Rogers and Nason (1998) use a structural VAR approach and
employ various identification schemes. They find Choleski-type identifications to yield
long-run dynamics that are inconsistent with long-run identification schemes in the spirit
of Blanchard and Quah (1989) and vice versa. They do however, not single out one
identification scheme that is superior to the others in its ability to identify global and
country-specific shocks. This would require cross-model evidence which we will provide
in this paper: the Choleski-identification scheme proposed in the previous section works
well in identifying global and country-specific shocks. We will argue that it focuses on an
immediate implication of the intertemporal approach (global shocks do not impinge on the
current account) whereas the Blanchard-Quah scheme will ensue in some intertemporal

30

models but not in others. After the model has been identified by the Choleski-scheme, it becomes possible to test the Blanchard-Quah scheme as an overidentifying restriction . We will now show that in the presence of a cointegrating relation it is particularly easy to test this overidentifying restriction.

Let for now the matrix $\mathbf{S} = \{s_{ij}\}_{i,j=1,2}$ define just any identification scheme such that $\mathbf{SS}' = \Omega$.

Then from $\boldsymbol{\varepsilon}_t = \mathbf{S}\mathbf{e}_t$ and $\eta_t = \boldsymbol{\alpha}'_\perp \boldsymbol{\varepsilon}_t$ we get

$$\eta_t = (\alpha_1 s_{21} - \alpha_2 s_{11})\, e_t^c + (\alpha_1 s_{22} - \alpha_2 s_{12}) e_t^w \tag{2.22}$$

Requiring that country-specific shocks be long-run neutral then amounts to

$$\frac{s_{21}}{s_{11}} = \frac{\alpha_2}{\alpha_1}$$

This is a testable proposition (conditional of course, on the identifying assumptions that give us $\mathbf{S}$): α_2 and α_1 are parameters of the reduced form and as such their estimates are unaffected by the identification scheme chosen. As shown e.g. in Johansen (1995), linear restrictions on the space spanned by $\boldsymbol{\alpha}$ can be tested and these tests are asymptotically χ^2-distributed. In the present setting, the hypothesis can be formulated as follows:

$$\boldsymbol{\alpha} = \mathbf{H}\psi \text{ where } \mathbf{H} = \begin{bmatrix} s_{11}/s_{21} \\ 1 \end{bmatrix}$$

If furthermore, we want to take account of the estimation uncertainty in s_{21}/s_{11}, this will no longer be a linear hypothesis on $\boldsymbol{\alpha}$ only. Still there is a simple way to test the hypothesis. Note that with $\Omega = \{\omega_{ij}\}_{j,i=1,2}$, for the Choleski-factor we have

$$\mathbf{S} = \begin{bmatrix} \sqrt{\omega_{11}} & 0 \\ \omega_{21}/\sqrt{\omega_{11}} & \sqrt{\omega_{22} - \omega_{21}^2/\omega_{11}} \end{bmatrix}$$

31

and hence $s_{21}/s_{11} = \omega_{21}/\omega_{11}$. Then in the framework of the conditional model

$$\Delta Y_t = \frac{\omega_{21}}{\omega_{11}}\Delta CA_t + \left(\alpha_2 - \frac{\omega_{21}}{\omega_{11}}\alpha_1\right) CA_{t-1} + lagged\ dynamics$$

testing the hypothesis we are interested in amounts to a $t-$test on whether the coefficient on CA_{t-1} is zero.

The Blanchard-Quah identification scheme links the period-zero impulse response of output and the current account, given by s_{21}/s_{11} to the relative long-run impulse response to (reduced-form) output and current account changes, given by α_2/α_1. This implies that the short-run dynamics of the system as given by the matrix $\mathbf{S}$ strongly influence the long-run dynamics and vice-versa. Under the Blanchard-Quah identification scheme, $\alpha_2 = 0$ implies $s_{21} = 0$ (note that in a cointegrated system $\boldsymbol{\alpha} = \mathbf{0}$ is not possible). Then, output is not only weakly exogenous in the long-run, but also, $\alpha_2 = s_{21} = 0$ implies that output is predetermined and also in the short-run unexpected output changes (which then coincide with global shocks) will drive the current account.

On the other hand, note that the Choleski-identification scheme we have suggested above will generically require the global shock to have some long-run impact on output: if $\mathbf{S}$ is the lower-Choleski-factor of Ω, $s_{12} = 0$ and $s_{22} > 0$. Hence, unless $\alpha_1 = 0$, i.e. we find the current account to be weakly exogenous, the Choleski-scheme will not be compatible with the Blanchard-Quah scheme w.r.t. to global shocks.

The preceding discussion puts us in a position to discuss the relative persistence of global and country-specific shocks. Recall the representation of the permanent shocks in (2.22) and note that the Choleski-identification scheme requires $s_{12} = 0$. Then

$$\eta_t = (\alpha_1 s_{21} - \alpha_2 s_{11}) e_t^c + \alpha_1 s_{22} e_t^w \tag{2.23}$$

The coefficient on e_t^c, $\alpha_1 s_{21} - \alpha_2 s_{11}$, is a function of the output- and current-account response in period zero: s_{21} measures the period-zero output response to a country-specific shock whereas s_{11} measures the corresponding current-account response. These

responses, in the long-run, get amplified by the coefficients α_1 and α_2. We can rewrite η_t as follows:

$$\eta_t = \alpha_1 \left[\left(\frac{s_{21}}{s_{11}} - \frac{\alpha_2}{\alpha_1} \right) s_{11} e_t^c + s_{22} e_t^w \right] \tag{2.24}$$

This equation tells us that the long-run impact of a one standard-deviation country-specific shock depends on the difference

$$\left(\frac{s_{21}}{s_{11}} - \frac{\alpha_2}{\alpha_1} \right)$$

The Blanchard-Quah identification scheme is compatible with the Choleski-scheme only if this difference is found to be zero.

The first ratio is the short-run impulse response of output relative to the current account. It tells us how a country-specific shock gets amplified in the period it occurs. The second term measures amplification as well, but now in the long-run: how much more strongly does output react to unexpected current account changes than to unexpected output changes?

Hence, we can interpret the difference between short-run and long-run adjustment as a measure of the relative contribution of country-specific shocks to the stochastic trend in output. Equivalently, we can understand it as a measure of the persistence of country-specific relative to global shocks. Because a measure of persistence should be positive, we here take the square of this difference and define:

$$\rho = \left(\frac{s_{21}}{s_{11}} - \frac{\alpha_2}{\alpha_1} \right)^2$$

Note also that this is a measure of persistence *net* of the relative variance of country-specific and global shocks: even if ρ is high, country-specific shocks may still explain a

33

small share of long-run variance because they are less volatile than global shocks. In this sense, ρ tells us how much more persistent country-specific shocks are than global shocks - regardless of their respective volatilities. We address this issue in the next subsection.

2.3.2 What drives the common trend?

The share of long-run output variance explained by country-specific shocks is given by

$$\frac{(\alpha_1 s_{21} - \alpha_2 s_{11})^2}{(\alpha_1 s_{21} - \alpha_2 s_{11})^2 + \alpha_1^2 s_{22}^2}$$

which from the previous section can also be written as

$$\frac{\rho s_{11}^2}{\rho s_{11}^2 + s_{22}^2}$$

If $\alpha_1 = 0$, then the country-specific shock will explain all trend output growth variance and ρ goes to infinity. Shocks to the current account (which are assumed to be country-specific) accumulate to the stochastic trend in output and there will be no long-run feedback from output to the current account. We can think of the economy being driven by idiosyncratic shocks that are transmitted from the rest of the world.

If, however, $\alpha_2 = 0$, then the shocks to output drive the joint dynamics of the system and the current account is the variable that has to bear the adjustment burden in the long-run. Still, the share of trend output variance in this case will not be zero but is given by:

$$\frac{s_{21}^2}{s_{21}^2 + s_{22}^2}$$

The relative weight of country-specific shocks will depend on the relative period-zero impulse response of output to global and country-specific shocks. So, country-specific shocks will still have their role but now we should think of them as originating in the country, with the output reaction causally prior to the reaction of the current account.

34

Econometrically, tests of the hypothesis $\alpha_{1,2} = 0$ amount to tests of weak exogeneity in the sense of Engle, Hendry and Richard (1983): the dynamics of the remaining variable in the system can be correctly captured by conditioning on the weakly exogenous variable in the sense that no long-run feedback relations are neglected. We present tests of this hypothesis in the empirical section of the paper.

2.3.3   Assessing the quality of shock identification

The identification of global and country-specific shocks in this model rests on insights derived from the theory: not all countries of the world can run current account surpluses or deficits simultaneously. Hence, the world interest rate should adjust and the effect on current accounts should be small or even zero.

Even though this seems a plausible assumption, it is clearly not testable in the framework of the model as the Choleski-decomposition we impose is just-identifying. However, our analysis will proceed in the same way for all major seven industrialized countries. Those countries account for roughly 60 percent of world economic output. How global or country-specific the shocks we identified actually are can be assessed using cross-country information. We will discuss this issue here.

A logical starting point is certainly to look at cross-country correlations of global and country-specific shocks. Here, we would expect that on average, global shocks are more highly correlated across countries than country-specific ones. But how far should we push this idea? It seems unlikely that cross-country correlations of country-specific shocks are actually zero - shocks might after all be specific to a group of countries. Also, some upward movements in the current account in one country will correspond to downward movements in another country's current account. This reflects transmission of shocks and the fact that when we use the current account as an identification device for asymmetric/country-specific shocks, this means that the shock does not have to originate in this country. Rather, the country-specific shock is the outcome of a country's lending to and borrowing from many other countries, essentially an amalgam of many bilateral

35

asymmetric shocks.

Likewise, global shocks should not be expected to be perfectly correlated. Rather, allowing for differences in internal transmission mechanisms, we should expect that the correlation is lower than unity.

An approach that takes account of the noise in the shock time series is principal component analysis. Let $\mathbf{E}_t^w = \{e_i^w\}_t^{i=1..7}$ be the vector of the stacked world-wide shocks and $\mathbf{E}_t^c$ be is the counterpart for the country-specific shocks. Then, the covariance matrix can be decomposed

$$\mathbf{cov}(\mathbf{E}) = \mathbf{P}\Lambda\mathbf{P}'$$
(2.25)

where $\Lambda = diag(\lambda_1....\lambda_7)$ and $\lambda i \geqq \lambda_{i+1}$ $i = 1..6$. The principal components are given by $\mathbf{P}'\mathbf{E}_t$, where the first principal component explains the highest share of the variance, the second the second-highest etc.

In particular, it becomes possible to test how many principal components are sufficient to explain the variation in the data. A test for this kind of problem has been suggested by Bartlett (1954). The hypothesis of the Bartlett test is that the first k principal components explain the variance of the data whereas the last $p - k$ (where p is the dimension of the vector $\mathbf{E}$) are essentially indistinguishable. For the determinant of the dispersion matrix of normalized variables (i.e. like the shocks we are dealing with) is

$$\det(cov(E)) = \prod_{i=1}^{p} \lambda_i$$

Furthermore, it is

$$trace(E) = \sum_{i=1}^{p} \lambda_i = p$$

36

Hence, under the null

$$\det(cov(E)) = \lambda_1 \lambda_2 ... \lambda_k \left\{ \frac{p - \sum_{i=1}^{k} \lambda_i}{p - k} \right\}^{p-k}$$

The alternative is that there are $k + 1$ significant principal components and the determinant of the dispersion matrix can then be written in an analogous way.

The ratio of the two determinants is given by

$$\left[\prod_{i=k+1}^{p} \lambda_i \right]^{-1} \left\{ \frac{\sum_{i=k+1}^{p} \lambda_i}{p - k} \right\}^{p-k} \tag{2.26}$$

When appropriately scaled with a factor involving sample size, the log of this expression can be given an approximate χ^2-distribution.

In the context of our problem, we would expect that such a test detects only one principal component that explains the variation in the data once we apply it to global shocks and a much larger number of significant principal components among the country-specific shocks.

Also, the theory suggests that the principal component driving the global shocks is associated with the world interest rate. We can test this implication by comparing $\mathbf{p}'\mathbf{E}_t^w$ with a measure of the world interest rate, where $\mathbf{p}'$ is the first row of $\mathbf{P}'$.

2.4 Empirical results

2.4.1 Estimation and model specifications

In this section, we report the results of the estimation of our model for the G7 countries. The data we used are annual real GDP from Gordon (1993), 1960-91 and current account / GDP ratios from Taylor (1996) and originally due to Obstfeld and Jones (1990). In order to make output volatilities comparable across countries, we transformed output into an index by dividing through by the first observation. We also divided the current

account by the first observation of output, i. e. we considered $\mathbf{X}_t = \begin{bmatrix} CA_t, & Y_t \end{bmatrix}'/Y_0$. Standard information criteria suggested that the seven models should be specified with one or two lags. We decided for two lags throughout. The model was then estimated with an unrestricted constant term.

We also included a number of conditioning variables in some of the models: in testing for the number of cointegrating relationships, we could not reject the null of no cointegration in the case of the US and Canada. This, however, should not be too surprising as the theoretical model is designed for a small open economy in that is treats the world interest rate as fixed. The US interest-rate, however,seems to play an important global role. Indeed, it is likely that the U.S.current account contains a large 'speculative' component that is the outcome of international capital flows induced by changes in the interest rate differential vis-a-vis the rest of the world..

We therefore decided to include the German-U.S. interest rate differential as an exogenous regressor into the model for the US. Even though we found the UK current account to be stationary, it is likely to be driven to a large extent by changes in the price of oil. Movements in the oil price, however, are prime candidates for global shocks, so we decided to condition the model for the UK on the price of oil.

In table 1 we present the results of Johansen's tests for cointegration after the inclusion of conditioning variables. Generally, we reject the null of no cointegration more strongly than without those variables. For six countries we find one cointegrating relationship at the 5-percent level. In particular we now also find a highly significant cointegrating relationship in the U.S. case. Only for Canada we continue to accept the null. Still we decided to impose one cointegrating relationship in the estimation of all seven models.

Table 1: Johansen's tests for cointegration

	Trace test		MaxEV test	
H_0	$h=0$	$h=1$	$h=0$	$h=1$
US	**30.35**	2.639	**27.71**	2.639
Japan	*17.04*	4.045	*13*	4.045
Germany	**18.2**	2.052	**16.15**	2.052
France	13.79	0.6392	*13.15*	0.6392
Italy	**25.68**	0.04728	**25.63**	0.04728
UK	**21.25**	4.096	**17.16**	4.096
Canada	10.25	0.4452	9.804	0.4452
90% crit. val	15.58	6.69	12.78	6.69
95% crit. val.	17.84	8.803	14.6	8.083

5 (10) %-significant values are in bold (italics)

Once we impose a cointegrating relationship in the estimation, tests of the cointegrating space show that it is generally the current account that is stationary: for six countries is the hypothesis that $\beta' = \begin{bmatrix} 1, & 0 \end{bmatrix}$ is accepted at the 5-percent level. For Germany there seems to be a small but significant coefficient on output in the cointegrating vector. Our unrestricted estimate of β for Germany is $\begin{bmatrix} 1, & -0.08 \end{bmatrix}$.

Table 2: tests on the cointegrating space $\beta' = \begin{bmatrix} 1, & \beta_2 \end{bmatrix}$

	US	Japan	Germany	France	Italy	UK	Canada
β_2	-0.0032	0.0107	-0.0848	0.0004	-0.0012	0.0543	0.0151
p-value	0.8302	0.4622	0.00106	0.941	0.8311	0.0927	0.253

Based on these pre-test results, we decided to proceed as follows: we imposed one cointegration relation in the estimation of all seven models. However, in the estimation of the German model we left the cointegrating space unrestricted.

39

2.4.2 Global and country-specific shocks

We are now in a position to discuss the quality of the identification scheme we have proposed for global and country-specific shocks.

We start by exposing the correlation matrices of global and country-specific shocks and their average value across countries (this cross-sectional mean excludes the country itself, of course) in table 3. Here, we find first favourable evidence that our scheme works fairly well. Global shocks are on average more highly correlated than country-specific shocks. Also, the p-values of the global shock are much lower and the cross-sectional mean is significant at conventional levels in four out of seven cases, whereas for the country-specific shock it is never found to be significant.

Table 3 a): cross country correlation of country-specific shocks

	US	Japan	Germany	France	Italy	UK	Canada
US	1						
Japan	-0.1932	1					
Germany	-0.2203	0.2888	1				
France	0.001465	0.2563	-0.1412	1			
Italy	-0.07919	0.2709	-0.06561	0.6595	1		
UK	0.09094	0.1825	-0.4724	0.1099	0.166	1	
Canada	0.1738	-0.2927	0.01252	-0.3498	-0.3039	0.03893	1
mean	-0.03773	0.08543	-0.09968	0.08936	0.1079	0.01932	-0.1202
std-dev.	0.1562	0.2588	0.2529	0.3484	0.3373	0.2464	0.2216
p-value	0.4094	0.3774	0.3549	0.4039	0.3809	0.4703	0.3054

values of cross-sectional means significant at 5 (10)% are in bold (italics)

40

Table 3 b): cross-country correlation of global shocks

	US	Germany	Japan	France	Italy	UK	Canada
US	1						
Germany	0.4021	1					
Japan	0.2999	0.283	1				
France	0.3714	0.4497	0.3642	1			
Italy	-0.07883	-0.116	0.3682	0.3681	1		
UK	0.2934	0.3706	0.4597	0.4495	0.203	1	
Canada	0.7015	0.161	0.3364	0.4039	0.255	0.5147	1
mean	0.3316	0.2584	**0.3519**	**0.4011**	0.1666	**0.3818**	**0.3954**
std-dev.	0.2507	0.2098	0.06279	0.04012	0.2147	0.1168	0.1929
p-value	0.1216	0.1364	0.00125	0.0000	0.2365	0.0111	0.0478

values of cross-sectional means significant at 5 (10)% are in bold (italics)

We then proceeded to test whether principal component analysis makes any sense in our setting. If shocks are spherical or at least independent, then there is no point in finding a rotation such that one direction explains as much as possible of the variance. In other words: orthogonalizing the variates would not carry any benefit in this case as the variates *are* already orthogonal. Before proceeding to an analysis of principal components, we therefore performed a test of independence for both E^c and E^w.

The test clearly rejected the null of independence for both types of shocks ($p-$values of 0.01 and 0.00). In the case of country-specific shocks, this suggests that international transmission of these shocks plays an important role.

Table 4 gives the results of the principal component analysis, subtable a) for the global shock and subtable b) for the country-specific shocks. The first principal component of the global shock identified for the G7 explains 43 percent of the variance whereas for the country-specific shock it accounts for only 30 percent of the variance. This hints at a higher degree of 'commonality' among the global shocks.

In the fourth column of the same table we also provide the results of the Bartlett tests

41

for dimensionality. At a conventional significance level of 5 percent, the tests suggests that country-specific shocks have one distinguishable principal components whereas the global shock displays five. This result seems somewhat at odds with our earlier finding that country-specific shocks have a lower cross-sectional correlation than global shocks. But note that once we lower the size of the test to 1 percent, then the principal components of the country-specific shock become indistinguishable whereas only two principal components survive for the global shock. Our results suggest that there is a reduced number of driving forces behind the global shocks. We will now try to identify these driving forces with observable economic variables. There are a few obvious candidates: as has been put forward in the introductory sections of this paper, theory suggests that changes in world interest rates are a prime candidate. Another variable is US-output growth.

Table 4 a): Principal component analysis of global shocks

Principal Comp.	Variance explained	Latent roots	Bartlett Test
1	43.66	3.056	2.981e-007
2	18.46	1.292	0.007342
3	13.48	0.9434	0.02079
4	9.463	0.6624	0.03481
5	8.208	0.5745	0.02402
6	4.612	0.3228	0.1096
7	2.12	0.1484	NaN

Table 4 b): Principal component analysis country-specific shocks

Principal Comp.	Variance explained	Latent Roots	Bartlett test
1	30.95	2.167	0.01094
2	23.54	1.648	0.05675
3	14.14	0.9901	0.2474
4	12.02	0.8413	0.1864
5	10.3	0.7211	0.1723
6	5.095	0.3566	0.7854
7	3.951	0.2766	NaN

The first and second principal components of the global shock are plotted in figure 1. Figure 2 gives the US output growth rate whereas figure 3 plots the US ex-post real interest rate.

Figure 4 plots the first principal component and the US output growth rate and figure 5 presents changes in the real interest rate and the second principal component.

The close comovement between US output growth and the first principal component that is apparent from the visual impression of figure 4 is confirmed by the correlation which is 0.68. There seems to be a link between the second principal component and the real interest rate but it does not show up very strongly in the correlation which is found to be 0.24. Also, this correlation is positive whereas from the theory we would expect that positive global shocks are associated with decreases in the real interest rate. Still, figure 5 suggests an important link between the two variables. We therefore proceeded to a more formal analysis of their joint time-series properties. Following the modelling approach suggested in Gonzalo and Granger (1995), we cumulated the second principal component of the global shock and the changes in the real interest rate. We then specified a cointegrated VAR in 2 lags:

$$\Gamma_z(\mathbf{L})\Delta\mathbf{Z}_t = \boldsymbol{\alpha}_z\boldsymbol{\beta}'_z\mathbf{Z}_{t-1} + \mathbf{v}_t$$

43

where $\mathbf{Z}'_t = \begin{bmatrix} \sum_{i=0}^{t} e_t^w, & r_t \end{bmatrix}$ and the covariance structure is given by

$$\Sigma = \mathbf{var}(\mathbf{v}_t) = \{\sigma_{ij}\}_{i,j=1,2}$$

We included an unrestricted constant and a step dummy to account for the secular increase in interest rates in the early eighties. Johansen's (1988) test suggested the presence of one cointegrating relationships. The estimated cointegrating vector was $\beta'_Z = \begin{bmatrix} 1, & 0.62 \end{bmatrix}$ and the hypothesis $H_0 : \beta'_Z = \begin{bmatrix} 1, & 1 \end{bmatrix}$ was accepted with p-value 0.2. This suggests that in the long-run changes in the real interest rate are perfectly inversely correlated with global shocks.

Tests also suggested that the real interest rate represents the common stochastic trend in $\mathbf{Z}_t$, i.e. we found $\alpha_{2Z} = 0$ which suggests that we can write a conditional model of the global shock:

$$e_t^w = \frac{\sigma_{21}}{\sigma_{22}}\Delta r_t + \left(\alpha_{1Z} - \frac{\sigma_{21}}{\sigma_{22}}\alpha_{2Z}\right)\left(\sum_{i=0}^{t-1} e_i^w + r_{t-1}\right) + lagged\ dynamics$$

Our estimate of σ_{21}/σ_{22} is -0.48, much higher in absolute terms than the correlation between Δr_t and e_t^w that we calculated earlier and that we found to be 0.24. Also, the correlation is now negative, in accordance with the theory.

The results suggest that the global shock is indeed negatively related to movements in the real interest rate. In the long-run the correlation seems perfect, whereas in the short-run it is somewhat less pronounced.

2.4.3   Persistence and the relative importance of global and country-specific shocks

Table 6 provides our estimates of persistence for country-specific shocks. The results are very interesting: for the four smallest economies, country-specific shocks are found to be much less persistent than global shocks,whereas for the G3, the U.S., Japan and Ger-

44

many, we find them to be 6-15 times more persistent than global shocks. This result may be due to two reasons: the G3 economies are large vis-a-vis the other four economies and therefore may find it difficult to fully smooth country-specific shocks through international borrowing and lending. Country-specific shocks may therefore become very persistent relative to global shocks. On the other hand, our procedure may suffer from some mismeasurement. As our results have shown so far, it is more likely to work well with a small open economy and country-specific U.S.-shocks are correlated with global shocks.

Table 6: Relative persistence of e^c vs. e^w

	US	Japan	Germany	France	Italy	UK	Canada
ρ	9.155	6.335	15.42	0.1721	0.7241	0.01657	0.026

In table 7 we test the overidentifying restriction imposed by the Blanchard-Quah identification, i.e. that $\rho = 0$. The first row in the table pertains to the 'naive' test in which we assume s_{11}/s_{21} fixed and just test a linear restriction on $\boldsymbol{\alpha}$. The second row gives the test based on the regression of ΔY_t on ΔCA_t, CA_{t-1} and lagged values. The 'naive' test clearly rejects the hypothesis for the US, Japan, Germany and Italy. This picture is not changing a lot once we do the regression test. However, the US becomes a borderline case now with the hypothesis accepted at the 13-percent level. In particular for the UK and Canada the data support the Blanchard-Quah identification. If we disregard the case of Italy, a general pattern is suggested by the data: the smaller the economy, the more likely are country-specific shocks to be long-run neutral w. r. t. output.

45

Table 7: Tests of the Blanchard-Quah restriction

Test on $\boldsymbol{\alpha}$

	US	Japan	Germany	France	Italy	UK	Canada
LR	13.44	9.067	15.2	0.923	15.41	1.069	0.4868
p-val.	0.0002	0.0026	0.0000	0.3367	0.0000	0.3011	0.4854

Regression test on $(\alpha_2 - \frac{\sigma_{21}}{\sigma_{11}}\alpha_1)$

	US	Japan	Germany	France	Italy	UK	Canada
t-val.	1.131	2.634	3.265	1.018	3.972	0.8713	0.1776
p-val.	0.134	0.0068	0.0014	0.1588	0.0002	0.1956	0.4302

LR is distributed as $\chi^2(1)$ and t-stat as $t(T-5)$ where $T = 32$ is the sample size

In table 8, we present the results of forecast error decompositions of changes in output and the current account. The result is interesting to contrast with our estimates of persistence: country-specific shocks seem to fully explain changes in the current account. This corroborates an important prediction of the intertemporal theory which predicts that the current account response to global shocks should be negligible. It also lends additional support to the validity of our identification scheme: if we think of a smooth current account response to structural shocks then we should not have done the data too much harm by imposing a zero-restriction in period zero.

It is interesting to compare the output decomposition with our estimates of the persistence of country-specific shocks: in the short-run global shocks explain the bulk of the output variance but the share of the country-specific shock is never negligible.

In the long-run the share of the country-specific shock increases, in particular so in the case of the G3. This reflects the high persistence of country-specific shocks in these countries. But note that even at the 10-year forecast horizon, country-specific shocks never explain much more than 50 percent of changes in output whereas the shocks where found to be 6-15 times as persistent as global shocks.

46

Table 8a: Variance share of ΔCA explained by country-specific shock

Forecast-horizon	US	Japan	Germany	France	Italy	UK	Canada
1	1	1	1	1	1	1	1
2	0.95	0.99	0.97	0.98	0.96	0.73	0.99
5	0.95	0.99	0.97	0.97	0.95	0.76	0.99
10	0.95	0.99	0.97	0.97	0.95	0.76	0.99

Table 8b: Variance share of ΔY explained by country-specific shock

Forecast horizon	US	Japan	Germany	France	Italy	UK	Canada
1 year	0.38	0.0000	0.00	0.03	0.27	0.31	0.12
2	0.50	0.04	0.01	0.14	0.61	0.31	0.11
5	0.52	0.31	0.48	0.15	0.67	0.38	0.13
10	0.53	0.34	0.57	0.15	0.67	0.38	0.13

Table 9 gives the share of trend output variance that is explained by country-specific shocks. In line with our earlier finding that country-specific shocks are very persistent in the G3 countries, the share of variance that can be ascribed to these shocks is between 20 and 30 percent for Japan and Germany and amounts to roughly 80 percent for the US. Among the smaller G7-economies, Italy is special in the sense that 40 percent of trend output variance is explained by the country-specific shock. For all other countries, the share of trend output variance explained by the country-specific shock is negligible.

Overall, the variance decompositions suggest that country-specific shocks are generally less volatile than global ones. The diagonal entries of **S** measure the variance of the structural shocks. Indeed, table 10 that gives the estimates of the ratio s_{11}/s_{22} shows that global shocks are generally one and a half (0.63^{-1})times as volatile as country-specific ones.

Table 9: Share of e^c in trend output variance

US	Japan	Germany	France	Italy	UK	Canada
0.80	0.20	0.29	0.14	0.41	0.00	0.01

47

Table 10: Relative variance of e^c and e^w- estimates of s_{11}/s_{22}.

US	Japan	Germany	France	Italy	UK	Canada	Average
0.3019	0.5039	0.6351	0.6456	1.288	0.5913	0.4634	0.6348

Table 11 provides the results of the tests for weak exogeneity, i.e. of the hypotheses $\alpha_i = 0$, $i = 1, 2$. It is interesting to note that with the exception of Italy we find that at the 5-percent level at least one variable is found to be weakly exogenous for all countries.

Table 11: Tests of weak exogeneity (p-values)

	US	Japan	Germany	France	Italy	UK	Canada
CA	0.00	0.01	0.13	0.00	0.00	0.00	0.00
Y	0.62	0.08	0.00	0.13	0.00	0.16	0.53

In the US and German cases, the current account is clearly found to be weakly exogenous. Note that, under the Choleski-identification, this amounts to saying that global shocks have no long-run effect on output. In both the German and US cases, the Blanchard-Quah restriction was found to be strongly rejected (table 7).

This is compatible with the picture that emerged earlier in which the U.S. output trend is purely domestically determined but acts as a generator for world-wide macroeconomic fluctuations. For Germany, the finding that the current account drives the common trend and the fact that a non-trivial cointegrating relationship prevails between output and the current account suggests that German trend output growth in the period 1960-91 has largely been driven by shocks to the export sector, a notion that is frequently referred to as 'export-led' growth. (see e.g. the study by Marin (1992))

2.5 Conclusion

In this paper, we have suggested using the reduced form of a simple intertemporal model of the current account to measure stylized facts in the international transmission of macroeconomic disturbances. We have proposed a simple identification scheme for global

48

and country-specific shocks. The identification scheme was assessed using cross-country evidence and seems to work reasonably well: global shocks are more highly correlated across countries than are country-specific shocks. Also, there are two dominant principal components among global shocks. Whereas one of them can straightforwardly be associated with US-output growth, the second one displays some short-run and perfect long-run correlation with a measure of the ex-post US real interest rate.

We have then used the proposed framework to collect stylized facts about the external adjustment of the G7 economies. Our results can be summarized as follows:

- Country-specific shocks account for most of the current account variance. This finding corroborates an important prediction of the intertemporal approach to the current account which suggests that the current account should react to the country-specific shock only.

- Country-specific shocks are much more persistent than global ones in the G3 economies and much less than global ones in the smaller G7 countries. Generally, the smaller the country, the less persistent are country-specific shocks.

- Country-specific shocks are generally found to explain only a moderate share of trend output growth.

- On average, global shocks are one and a half times more volatile than country-specific ones.

- Global shocks have two dominant principal components: the more important one is found to be highly correlated with US output growth. In accordance with the intertemporal approach to the current account, the second one is in the long-run perfectly negatively correlated with the real interest rate. In the short-run there seems to prevail a smaller negative correlation.

- Changes in the US interest rate seem to trigger important current account reactions that are then found to be statistically exogenous w.r.t. to output dynamics in this

49

country.

- In Germany, there is a non-trivial cointegrating relationship between output and the current account. Also, the current account seems to drive the stochastic trend in output as it is found to be weakly exogenous. Evidence for the German case seems inconclusive. We propose to interpret our findings as evidence of Germany's output growth over the period being driven by export-shocks.

50

Bibliography

[1] Barro, E. and Sala-i-Martin, X (1990),'World interest rates' in O. Blanchard and S. Fischer, eds., *NBER macroeconomics annual* 1990.

[2] Bartlett, M. S. (1954), 'A note on the multiplying factors for various χ^2-approximations', *Journal of the Royal Statistical Society*, Ser. B, 16, 296.

[3] Bayoumi, Tamim and Taylor, Mark P., (1995), "Macroeconomic Shocks, the ERM and Tripolarity," *The Review of Economics and Statistics*, 321-331.

[4] Bayoumi, Tamin and Eichengreen, Barry, (1992a), 'Shocking Aspects of European Monetary Union,' CEPR Discussion-Paper no. 643.

[5] Bayoumi, Tamin and Eichengreen, Barry, (1992b), 'Macroeconomic Adjustment under Bretton Woods and the Post-Bretton-Woods Float: An Impulse Response Analysis,' CEPR Discussion Paper no. 729.

[6] Beveridge, Stephen and Nelson, Charles R., "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the 'Business Cycle"', *Journal of Monetary Economics* 7, 151-173.

[7] Blanchard, Olivier and Quah, Danny, (1989), "The Dynamic Effects of Aggregate Demand and Supply Disturbances," *American Economic Review* 79, 655-73.

[8] Campbell, J. Y, and Shiller, R.J., (1987), ' Cointegration and tests of present value models,' *Journal of Political Economy* 95: 1062-1088.

[9] Engle, R.F., Hendry, D.F. and Richard, J.-F. (1983), 'Exogeneity', *Econometrica* 51: 277-304.

[10] Engle, Robert F. and Granger, Clive W. J. (1987), "Cointegration and Error-Correction: Representation, Estimation and Testing," *Econometrica* 55: 251-76.

[11] Gonzalo, Jesus and Granger, Clive, (1995), "Estimation of Common Long-Memory Components in Cointegrated Systems," *Journal of Business and Economics Statistics* 13: 27-35.

[12] Gordon, R.J. (1993), *Macroeconomics*, Addison-Wesley. Data set also to be found at *http://www.nber.org/pub/gordon*.

[13] Johansen, Søren, (1988), "Estimation Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", *Econometrica* 59: 1551-80.

[14] Johansen, Søren, (1995), *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press.

[15] Johansen, Søren, (1997), "Common Trends in the Cointegrated VAR Model," *European University Institute*, mimeo.

[16] Kendall, M. and Stuart, A. (1975), *The Advanced Theory of Statistics*, vol. 3. London.

[17] King, Robert, Plosser, Charles I., Stock, James H. and Watson, Mark W., (1991), "Stochastic Trends and Economic Fluctuations", *American Economic Review* 81: 819-40.

[18] Marin, D. (1992), 'Is the Export-Led Growth Hypothesis valid for industrialized countries?', *Review of Economics and Statistics* 74: 678-88.

[19] Nason, J. and Rogers, J. H. (1998), 'Investment and the Current-Account in the Short-Run and the Long Run', mimeo, University of British Columbia.

[20] Obstfeld, M., (1986), 'Capital Mobility in the World Economy: theory and measurement', *Carnegie Rochester Conference Series on Public Policy* 24: 55-104.

[21] Obstfeld, M., (1995), 'International Capital Mobility in the 1990s' in Kenen, P. (ed.) *Understanding Interdependence*, Princeton University Press, 1995.

[22] Obstfeld, M. and Rogoff, K. (1995), 'The intertemporal approach to the current account,' NBER working paper 4893.

[23] Obstfeld, M. and Rogoff, K., (1996), *Foundations of International Macroeconomics*, MIT Press, Cambridge.

[24] Sachs, J, (1981), 'The Current Account and Macroeconomics adjustment in the 1970s,' *Brookings Papers on Economic Activity* 1: 201-268.

[25] Stock, James H. and Watson, Mark W., (1988), "Testing for Common Trends", *Journal of the American Statistical Association* 83, 1097-1107.

2.6 Figures
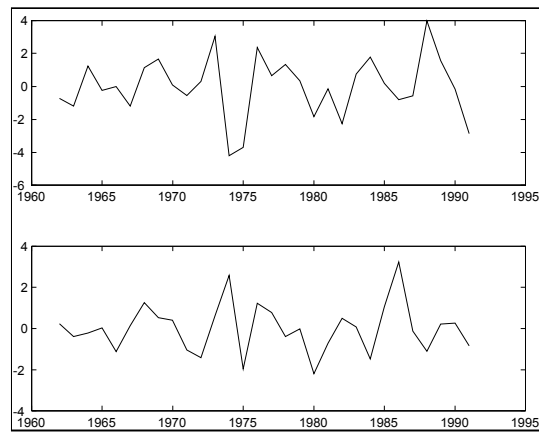
Figure 2-1: First (upper panel) and second principal components of the global shocks
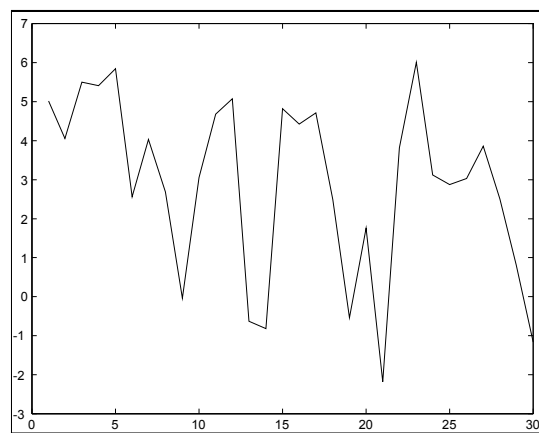
Figure 2-2: US GDP growth rates 1960-91.

54

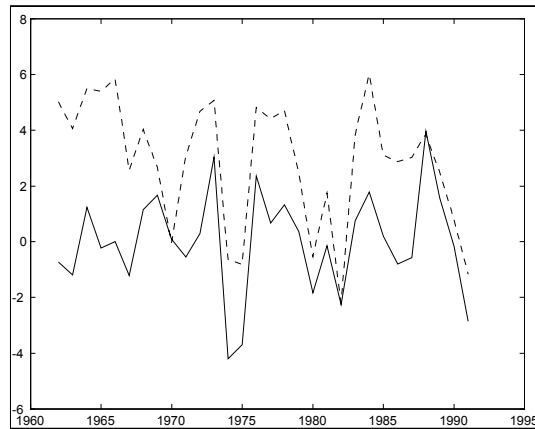Figure 2-3: US real interest rate (ex-post, based on GDP-Defl.)

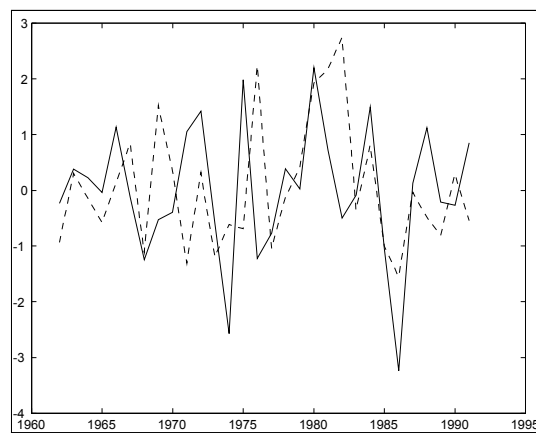Figure 2-4: US GDP growth rates and the first principal component of global shocks

55

Figure 2-5: Changes in the real interest rate (dashed) and second principal component of global shocks

56

Chapter 3

The Feldstein-Horioka Puzzle and a New Measure of International Capital Mobility

3.1   Introduction

In a world with perfect capital mobility, a country can always run current account deficits if its desire to consume and invest cannot be funded domestically. This basic insight provided the motivation for the seminal paper by Feldstein and Horioka (1980) in which the authors found very high savings-investment correlations for a large cross-section of countries. Their result has long been perceived as a puzzle and constitutes a challenge to the view that world capital markets are well integrated. In the presence of perfect capital mobility, investment should go where it yields the highest real returns, whilst consumption should depend only on the permanent value of income, not on contemporaneous investment decisions.

Subsequent research has rationalized the comovement of domestic saving and investment even in the presence of perfect capital mobility. Obstfeld (1986, 1995) and Obstfeld and Rogoff (1995) have pointed to two possible mechanisms that can generate the ob-

served correlation. In a small open economy, total factor productivity shocks that are sufficiently persistent can create positively correlated impulse responses of savings and investment. This mechanism is also suggested in Mendoza (1991). The second mechanism relies on global shocks that impinge on both savings and investment simultaneously. This is the channel formally explored in Baxter and Crucini (1993).As Coakley, Kulasi and Smith (1998) point out, the consequence of these theoretical results was that it has become a consensus in the profession that savings-investment correlations are not very informative about capital mobility.

In the present paper, we provide further justification for this view but contrary to the aforementioned rationalizations it is based on the reduced-form implications of the intertemporal approach to the current account. Hence, it does not have to rely on structural assumptions about the kind of shocks that are hitting an economy. We find that any correlation between savings and investment can ensue in a simple model of current account behaviour with perfect capital mobility and that under reasonable assumptions this correlation can be close to unity. Yet, the spirit of the Feldstein-Horioka approach, namely that inference on international capital mobility is possible from savings and investment data alone, can be preserved.

Under the assumptions of the theory and the additional assumption that the macroeconomic aggregates savings investment and output are very persistent, non-stationary processes , the joint dynamics of savings and investment is appropriately specified in the form of a vector error-correction model (VECM). This econometric specification allows to distinguish clearly between short-run and long-run capital mobility. The measure of short-run capital mobility is a suitably adjusted correlation, similar to the one suggested by Feldstein and Horioka, whereas the measure of long-run international capital mobility (ICM) is based on Johansen's (1988) procedure for estimating the cointegrating space.

The original work by Feldstein and Horioka (1980) emphasised the high correlation of savings and investment in a cross-section, whereas formal theoretical rationalizations of the correlation - like the ones mentioned before - mainly aim at explaining the time series

58

behaviour of the two variables. Also in the present paper, the analysis will be confined to the time series properties of savings and investment[1].

It is not within the scope of this paper to attempt to survey the huge literature on the Feldstein-Horioka finding (for a recent survey see Coakley, Kulasi and Smith (1998) or Obstfeld and Rogoff (1995)). There is, however, a recent trend towards vectorautoregressive and cointegration methods to address the topic. As this paper makes use of these techniques, we will briefly summarize some of this research:

Ghosh (1995) has used an intertemporal model to derive a desired current account from observed data. He finds that the desired current account tracks the actual current account reasonably well, hence providing evidence in favor of perfect capital mobility.

Moreno (1997) has suggested to interpret the degree of short-run divergence in the impulse responses of savings and investment as a measure of capital mobility.

Taylor and Sarno (1997) used the structural VAR approach pioneered by Blanchard and Quah (1989) to decompose savings and investment into permanent and transitory components. They find that transitory components of UK/US savings and investment are more highly correlated than changes in the permanent components. They claim that this finding is consistent with the presence of frictions in international capital markets. Only if innovations are permanent does investment flow abroad and the link between savings and investment is loosened. If, however, shocks are transitory, then the cost of investing abroad might be too high due to market frictions and a high correlation between saving and investment comes about. However, their results are supportive of the notion that capital mobility has increased in the 1980s: they report short-run correlations between savings and investment for the period 1979-1994 that are significantly lower than for the 1955-1979 period.

The remainder of the paper is organized as follows: section 2 presents a simple model of current account dynamics based on intertemporal optimization. These models were

---

[1]It should be noted however, that a time series-rationalization is in some way more fundamental: if savings and investment move one to one over time in an individual economy and do so for all economies under study, then, of course, the cross-section correlation will be trivially unity as well.

59

first applied to current account dynamics by Sachs (1981). Section 3 discusses the classical Feldstein-Horioka regression. We demonstrate that any correlation between the transitory parts of savings and investment can ensue and that these correlations *per se* do not contain any information about capital mobility. In Section 4 we suggest a new measure of long-run international capital mobility (ICM) which is easily calculated as a by-product of Johansen's (1988) procedure for the estimation of the cointegrating space. Section 5 applies our insights to a unique set of long-run historical data from the United Kingdom and the United States. Section 6 concludes.

3.2   Current account models and cointegration

This section examines the implications of the intertemporal model of the current account in the spirit of the work by Sachs (1981) or as discussed in Obstfeld and Rogoff (1995). We use a simple variant of the model which considers a small open economy where the world interest rate is fixed at r and utility is quadratic in consumption. In such a model, the current account can be represented as the discounted sum of expected changes in net output:

$$CA_t = -\sum_{i=1}^{\infty} R^{-i} E_t(\Delta NO_{t+i}) \tag{3.1}$$

Here, $R = 1 + r$ and net output is defined as gross national product minus government consumption and investment:

$$NO_t = Y_t - I_t - G_t \tag{3.2}$$

60

The current account itself is defined as the difference between savings and investment:

$$CA_t \equiv S_t - I_t \tag{3.3}$$

The present-value relationship (1) together with the definition (3) defines a cointegrating relationship that is typical of present-value models: If net output, saving and investment can be characterized as $I(1)$-processes, then ΔNO_t will be $I(0)$ and so will be CA_t as the discounted sum of ΔNO_t. Hence, saving and investment cointegrate with cointegrating vector

$$\boldsymbol{\beta} = \begin{bmatrix} 1, & -1 \end{bmatrix}' \tag{3.4}$$

This result of current account stationarity is very robust with respect to the specification of the intertemporal model. In particular, the assumptions made above about quadratic utility and a fixed world interest rate can be relaxed. As Obstfeld (1995) has discussed, present-value relationships like (3.1) will arise in much more complicated and richer models. In particular, it is likely to survive in a model setup where there are barriers to capital mobility; the nation's budget constraint has to be respected no matter how mobile or immobile capital is.

3.3   The Feldstein-Horioka regression

In their seminal paper, Feldstein and Horioka (1980) performed a regression of the form

$$i_t = a + bs_t + u_t \tag{3.5}$$

where lower case letters denote variables as shares of GDP, i.e.$i = I/Y$ and $s = S/Y$. We will refer to (3.5) as the "classical" FH regression and the FH puzzle has generally been expressed in terms of estimates of b that are found to be close to one.

61

We will now look at two notions of correlation between savings and investment: between the levels and between the suitably extracted transitory components. Throughout the remainder of the paper, we will deal with savings and investment *rates*, even though we will at times leisurely refer to i and s as 'investment' and 'savings'.

Suppose, i_t and s_t can be characterized as $I(1)$-processes. As investment and saving cointegrate with cointegrating vector $[1, -1]$, there will be an error-correction representation of the form[2]:

$$\Gamma(\mathbf{L})\Delta \begin{bmatrix} s_t \\ i_t \end{bmatrix} = \boldsymbol{\alpha} CA_{t-1} + \boldsymbol{\varepsilon}_t = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} CA_{t-1} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \qquad (3.6)$$

where $\Gamma(L) = I - \sum_{i=1}^{k} \Gamma_i L^i$ is a 2×2–matrix polynomial in the lag-operator L , ε_{1t} and ε_{2t} are white-noise disturbances and Δ is the difference operator.

The cointegrating relationship imposes a long-run one-to-one relationship between investment and saving. Define the permanent value of a stochastic-process X_t as today's value plus the sum of all forecastable changes:

$$\mathbf{X}_t^P = \mathbf{X}_t + \sum_{i=1}^{\infty} \mathbf{E}_t(\Delta \mathbf{X}_{t+i}) \qquad (3.7)$$

This definition of a permanent value naturally leads to the Beveridge-Nelson (1981) decomposition (see also Proietti (1997)). Because a country can not permanently invest more or less than it saves, the permanent value of savings and investment have to move together one for one:

$$i_t^P = s_t^P \qquad (3.8)$$

We also derive this result formally in appendix A.

---

[2]We will ignore the constant term in our theoretical derivations.

62

Hence, the typical Feldstein-Horioka regression of investment on saving rates is just a cointegrating regression in the sense of Engle and Granger (1987). The OLS-estimator is known to be superconsistent in this case and a regression coefficient of unity just reflects the long-run relationship between savings and investment.

Another notion of correlation in this context refers to the comovement of the stationary part of the series after appropriate detrending: how should we expect $i_t - i_t^P$ and $s_t - s_t^P$ to correlate and what can we learn from the correlation of the transitory components?

In appendix B we derive the following expression

$$\left[\begin{array}{c} s_t - s_t^P \\ i_t - i_t^P \end{array} \right] = \mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t = \boldsymbol{\psi}CA_t + \boldsymbol{\beta}_\perp f_t \tag{3.9}$$

where $\mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t$ is the cyclical component of the Beveridge-Nelson decomposition, $\boldsymbol{\beta}'_\perp = \left[\begin{array}{cc} 1, & 1 \end{array} \right]$ is the orthogonal complement of $\boldsymbol{\beta}$, f_t is a univariate stationary stochastic process and $\boldsymbol{\psi}' = \left[\begin{array}{cc} \psi_1, & \psi_2 \end{array} \right]$ a two-dimensional vector.

Equation (3.9) states that $\mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t$ can be decomposed into one part which captures the error correction of the model, $\boldsymbol{\psi}CA_t$, and another part, given by $\boldsymbol{\beta}_\perp f_t$ which is pure short-run dynamics. Note that $\boldsymbol{\beta}'_\perp = \left[\begin{array}{cc} 1, & 1 \end{array} \right]'$ and therefore the pure short-run dynamics of savings and investment are perfectly positively correlated. Also the error correction dynamics are perfectly correlated but either positively ($\psi_1 \psi_2 > 0$) or negatively ($\psi_1 \psi_2 < 0$). The variance of $\mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t$ is given by

$$Var(\mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t) = \boldsymbol{\psi}\boldsymbol{\psi}'\sigma_{ca} + \left(\boldsymbol{\psi}\boldsymbol{\beta}'_\perp + \boldsymbol{\beta}_\perp \boldsymbol{\psi}' \right)\sigma_{f,ca} + \boldsymbol{\beta}_\perp \boldsymbol{\beta}'_\perp \sigma_{ff}$$

where σ_{ca} and σ_{ff} denote the variances of the current account and the common factor f_t respectively and $\sigma_{f,ca}$ is the covariance between the two. The correlation between the

63

components of $\mathbf{C}^*(\mathbf{L})\varepsilon_t$ is then given by:

$$
\begin{aligned}
\rho \quad &= corr(\mathbf{e}_1' C^*(L)\varepsilon_t, \mathbf{e}_2' C^*(L)\varepsilon_t) \\
&= \frac{\psi_1\psi_2\sigma_{ca}+(\psi_1+\psi_2)\sigma_{f,ca}+\sigma_{ff}}{\left[\left(\psi_1^2\sigma_{ca}+2\psi_1\sigma_{f,ca}+\sigma_{ff}\right)\left(\psi_2^2\sigma_{ca}+2\psi_2\sigma_{f,ca}+\sigma_{ff}\right)\right]^{\frac{1}{2}}}
\end{aligned}
\tag{3.10}
$$

Here, $\mathbf{e}_1$ and $\mathbf{e}_2$ are the first and second unit vectors.

In general, this expression will depend on the variance-covariance structure of CA_t and f_t but also on ψ_1 and ψ_2.

In the stationary case, the Feldstein-Horioka approach predicts that under high capital mobility the variance of the current account should be high relative to the variance of savings and investment. Because

$$
Var(CA_t) = Var(s_t) - 2Cov(s_t, i_t) + Var(i_t)
\tag{3.11}
$$

we have

$$
Cov(s_t, i_t) = \frac{1}{2}\left(Var(s_t) + Var(i_t) - Var(CA_t)\right)
\tag{3.12}
$$

Hence, a low savings-investment correlation requires that

$$
\frac{Var(CA_t)}{Var(s_t) + Var(i_t)}
\tag{3.13}
$$

be near unity. This insight, however, does not carry over to the non-stationary case because the unconditional second moments of s and i will not exist. Still, if the variability of the current account buffers a large share of the variance in the transitory part of savings and investment, this can be interpreted as indication of high capital mobility. But note that this is not equivalent to a high correlation of the stationary components of savings and investment: even if error-correction explains all the variance of the transitory

64

dynamics, i.e.

$$\frac{\psi_1^2 \sigma_{CA}}{Var(s_t - s_t^P)} = \frac{\psi_2^2 \sigma_{CA}}{Var(i_t - i_t^p)} = 1 \tag{3.14}$$

(which in turn implies $\sigma_{ff} = \sigma_{f,ca} = 0$), the correlation coefficient ρ can be plus or minus unity, depending on the sign of $\psi_1 \psi_2$. If savings and investment are non-stationary, error-correction behaviour, embodied in the coefficients ψ_1 and ψ_2, is likely to obscure the informational content of savings-investment correlations with respect to capital mobility, even after the variables have been rendered stationary. In the next section, we address the issue whether savings and investment data contain information about capital mobility at all. We are going to argue that it is just the error-correction itself that is interesting.

3.4 Inference on international capital mobility using savings and investment data

In this section, we will argue that the essence of the Feldstein-and-Horioka argument can be saved: inference on capital mobility is possible from saving and investment data alone.

To illustrate our notion of long-run capital mobility, consider the case of current account targeting discussed in Artis and Bayoumi (1992). Past current account deficits might incur government action in the sense that the government tends to offset private sector behaviour by increasing public sector savings or by trying to induce the private sector to increase its savings through policy action such as capital controls or monetary policy measures such as higher interest rates. No matter what the details of government action look like, however, in these circumstances one would probably expect a stronger predictive power of past current account (levels) for today's movements in national savings.

To measure capital mobility, we suggest to look at the adjustment coefficients in the bivariate VECM representation of our savings-investment system, i.e. at $\alpha = \begin{bmatrix} \alpha_1 & \alpha_2 \end{bmatrix}$.

65

Suppose α_1, is close to zero. In this case, past current accounts have only a small impact on present changes in savings, i.e. today's savings decision is relatively independent of the budget constraint and hence savings and investment become dichotomous in the sense implied by Feldstein and Horioka. Conversely, a small absolute value of α_2 indicates that domestic investment opportunities can be exploited, regardless of what the current account, i.e. the country's past savings and investment decisions used to be.

While the information we could gain by looking at α_1 and α_2 separately is certainly valuable, the focus in the literature on univariate modelling can also be explained in terms of the desire to have a composite measure of capital mobility. We will therefore suggest a measure of long-run capital mobility that arises naturally as a function of the parameters of our reduced-form model.

Johansen (1988), (1996) has shown that the estimation of the cointegrating space in a VECM is essentially a generalized eigenvalue problem. The maximum eigenvalue ensuing from the solution of this problem can be given the representation

$$\Lambda = \widehat{\alpha}' \widehat{\Sigma}_{00}^{-1} \widehat{\alpha} \tag{3.15}$$

where $\widehat{\Sigma}_{00}$ is the estimate of the variance-covariance structure of the first auxiliary regression in the Johansen (1988) procedure. The asymptotic distribution of Λ and procedures for the estimation of its covariance have recently been worked out by Hansen and Johansen (1998).

The nice property of Λ is that it is always between zero and one. Our argument here is that a high value of Λ implies low capital mobility whereas a low value of Λ is tantamount to a high level of capital mobility. Note in particular, that once Λ is zero this implies that the system has two cointegrating relationships, hence s and i are difference stationary but do not cointegrate. But this is exactly what we meant to imply previously: under perfect capital mobility, the system should still revert to equilibrium, i.e. cointegration and error correction should be present but should not be very strong. And this just implies a small (but significant) Λ.

66

Let us relate our indicators of long-run capital mobility to others suggested in the literature:

In a recent paper, Feldstein and Bacchetta (1991) estimated a specification of the form

$$i_t = a + b(i_t - s_t) + u_t \tag{3.16}$$

thus modifying the classical FH regression to allow for some kind of long-run equilibrium adjustment. As Taylor (1996) pointed out, if i and s are non-stationary but cointegrate, (3.16) will be misspecified. He suggested to estimate a univariate error correction model (ECM)

$$\Delta i_t = a^{ECM} + b^{ECM} \Delta s_t + c^{ECM}(s_t - i_t) + v_t \tag{3.17}$$

He then proposed to interpret the coefficient b^{ECM} as a measure of short-run capital mobility and c^{ECM} as a measure of long-run capital mobility. This line of reasoning is very close to ours. Notice, however, that in terms of the parameters of the VECM, Taylor's regression can be interpreted as a conditional model of investment, given savings. Conditioning investment on savings yields

$$\Delta i_t = \omega \Delta s_t + (\alpha_2 - \omega \alpha_1) ca_{t-1} + \textit{lagged dynamics} \tag{3.18}$$

where ω is a linear function of the covariance structure of the reduced form errors given by

$$\omega = \Omega_{12}\Omega_{11}^{-1} \text{ and } \Omega = \begin{bmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{bmatrix} = E(\varepsilon_t \varepsilon_t')$$

The coefficient ω measures short-run capital mobility - it is often referred to as a short-run savings retention coefficient (Taylor (1996)). It is a function of the covariance

67

of the reduced form errors, i.e. those innovations in savings and investment that are unexplained by our model. And as such, for once, a high value of ω is nothing that we should expect from the theory. Hence, low values of ω can be interpreted as indicative of high short-run capital mobility: changes in savings do not have high predictive power for contemporaneous changes in investment.

In as far as ω is interpreted as measure of short-run capital mobility, note that the coefficient c^{ECM} from equation (3.17) is a function not only of both coefficients of $\boldsymbol{\alpha}$ but also of short-run capital mobility. Hence, c^{ECM} does generally not tell us anything about how sustainable a country's current account position actually is, and hence is informative about the true adjustment process only if $\alpha_1 = 0$.

The system approach we suggest in this paper, gives us two measures of international capital mobility: one, the short-run retention coefficient is nothing else than a regression of the reduced form errors of investment on those of savings and tells us how investment and savings are correlated *net* of the working of the intertemporal model. The other one, based on the generalized eigenvalue problem underlying the estimation of a cointegrated system, is a composite measure of how sustainable a country's current account position is and, as such, measures long-run mobility. To our knowledge, the literature has so far not exploited such a system-based approach to disentangle short-run and long-run capital mobility cleanly.

In the next section we apply our insights to a unique data set due to Taylor (1996).

3.5 Empirical Results

In this study we use a unique set of long-range annual data on national savings and investment rates compiled and first used by Taylor (1996) to study the topic of international capital mobility. Data for the United Kingdom range from 1850-1992, data for the United States is from 1874 to 1992. Figures 1 and 2 provide a plot of the data set for the two countries.

We first estimated an unrestricted VAR with a constant. Following the Schwarz-, Hannan-Quinn and Akaike criteria we specified the model with two lags. We performed Johansen's test for cointegration. The results, given in table 1, suggested one cointegrating relationship for the US whereas in the model for the UK, no cointegrating relationship was detected. Once we imposed two step dummies for WWI and WWII, however, we found cointegration also in the UK-model. Visual inspection of the data, suggests that there are a number of structural breaks, most notably the two world wars. Our cointegration tests might therefore be invalidated because of parameter non-constancy. Following our theoretical specification, we imposed one cointegrating relationship and then proceeded to estimate Λ recursively, following the procedure developed in Hansen and Johansen (1998): if the maximum eigenvalue vanishes, there will be no cointegration between the variables.

Figure 3 and 4 give the results of this recursive estimation for the UK and the US respectively. It becomes apparent that the parameters of the model are not stable over the sample period and that a secular break occurred during WWI. Strictly speaking, parameter estimates after the structural break are not valid. Still the graphs support the interpretation that long-run international capital mobility recovered quickly after WWI, soon reaching pre-war levels. Also neither the Great Depression nor WWII seem to have disrupted long-run international capital mobility very strongly.

For the United States, WWI seems to have been particularly disruptive to ICM. But our estimates suggest that long-run international capital mobility quickly increased after WWI and that already during the great depression, it reached pre-WWI levels. After that, international capital mobility for the U.S. seems to have remained more or less constant over the rest of the sample period, with no major disruptions during the second world war nor further marked increases in ICM in the Bretton Woods or post-Bretton Woods periods.

For the UK before WW-I, we find relatively low levels of long run capital mobility. The variance of the estimate is rather high, though, and indeed we cannot reject the

hypothesis of equality of long-run capital mobility before and after world war one. As in the US case, WWI has disrupted long-run capital mobility severely but in the UK the sustainability of the current account position recovers even quicker than in the United States and stays roughly constant for the rest of the sample period, with the exception of WWII where ICM seems to reach a new peak. We believe that this is due to the exceptional financial aid the UK received from the United States during WWII. Current account deficits have been large in that period but will not have triggered appropriate reactions in savings and investment rates. This will bias the estimates of α downwards.

In spite of high correlations between savings and investment, long-run capital mobility over the century seems to have been remarkably high - at least for the United States and the United Kingdom. The first world war seems to have been disruptive to long run capital mobility but both countries were able to recover long-run sustainable current account positions soon. Our findings suggest that the role of the great depression as a watershed for ICM, as suggested in Eichengreen (1990) and Taylor (1996), is not quite warranted for the two countries. The difference in our results vis-à-vis Eichengreen and Taylor might arise because our analysis so far has exclusively focussed on long-run capital flows. The formal setup of our model allows us to distinguish cleanly between the short and the long-run and it seems plausible that the great depression was less disruptive to long-run ICM than to short-run capital flows.

Given that our results for long-run capital mobility differ somewhat from those reported in the literature, we also set out to estimate short-run capital mobility. After all, our measure of long-run capital mobility is entirely free of short-run dynamics whereas the coefficient c^{ECM} in regression (3.17) is in fact a function of ω, as the conditional model in (3.18) shows. Figures 5 and 6 plot the estimate of the short-run savings retention coefficient. Here a break occurs during WWI but whereas in the United States short-run (SR) capital mobility recovers after the war, it remains low in the UK. In contrast to LR capital mobility, SR capital mobility seems to have suffered a further setback during the great depression and during WWII from which it did not recover after 1945. Rather, for

70

the UK, SR capital mobility tends to decline and only the demise of the Bretton Woods system seems to have brought it back to pre-WWII levels. For the US the demise of Bretton Woods does not seem to have influenced the savings-investment correlation.

For both the UK and the US, figures 5 and 6 suggest that there are four regimes governing short-run capital mobility:

- the pre-world war I period of the classical gold standard, 1880-1913. As Bayoumi (1990) has claimed this was the one historical period that came closest to the paradigm of perfect capital mobility.

- The interwar period, a period that Taylor (1996) and Obstfeld and Taylor (1996) have found to be one of secular barriers to capital mobility. Taylor (1996) has included the two world wars in this subsample. We do not follow him in this respect but rather restrict ourselves to the period 1919-39. During the two world wars, the US was giving immense financial and material aid to the UK. Hence, the UK was running huge current account deficits which were financed mainly from US current account surpluses. These huge and extraordinary government transfers are likely to bias downwards the estimates of α in our method, as neither the US nor the UK are likely to have been concerned with current account deficits in their wartime policymaking.

- The postwar period up to the breakdown of the Bretton Woods system, 1946-71.

- The post Bretton Woods period, 1971-92, stretching to the end of the sample.

This classification is in line with the one given in Taylor (1996).

We decided to estimate the model for the four subperiods identified above. This certainly poses small-sample problems but gives us the benefit of parameter constancy. Again, we imposed one cointegrating relationship in the estimation. Our estimates of Λ are given in table 4.

71

In line with our recursive estimates, for the UK we find the point estimate of Λ to be smaller in that period than in the pre-WWI era. However, given the size of the standard errors, we can not reject equality.

For the United States, our findings are consistent with the results of Eichengreen (1990) and Taylor (1996), who identified the interwar period as an era that was particularly disruptive to international capital mobility. What is surprising however, is that our results suggest that capital mobility continued to fall in the post WWII-era and this for both the UK and the United States. In the case of the United States, only the post-Bretton Woods period sees levels of ICM that are comparable with those of the classical gold standard.

However, the post-WWII subperiods are very short and figures 3 and 4 suggest that they are indeed not very heterogeneous. Also, the results for the immediate post-WWII (i.e. the Bretton Woods) period may be heavily influenced by the huge current account surpluses (deficits) that the US (UK) experienced in the immediate aftermath of the war.

In order to achieve a comparison of long-run capital mobility under the classical gold-standard with post-WWII capital mobility, we therefore merged the third and fourth subperiods, dropping the immediate aftermath of the war, i.e. we used data from 1950 to the end of the sample, 1992. To account for potential parameter instability, we included a step-dummy for the post-Bretton-Woods period and the oil price for the UK which became an oil exporter in the late seventies.

Information criteria suggested three lags for the two models and we estimated both with an unrestricted constant. One cointegrating relationship was found at the 10-percent level in both cases, the maximum eigenvalue test even indicated cointegration at the 5 percent level for the UK (table 5). We then tested restrictions on the cointegrating space and the hypothesis that β' equals its theoretical value $\begin{bmatrix} 1 & -1 \end{bmatrix}'$ was accepted at high significance levels (table 6).

For Λ we found values of 0.30 for the UK and 0.19 for the US. Table 7 gives the results with standard errors. The point estimates for Λ in the post-WWII period are lower than

72

the ones for the period of the classical gold standard (table 4, line 1). Taking account of the 95-percent confidence intervals, we cannot reject the hypothesis that they are equal, though.

Our findings suggest that the transition from the Bretton-Woods system to floating exchange rates has had very little impact on long-run international capital flows. Also for short-run capital mobility,according to our recursive estimates, the effects for the two countries in our study were moderate. Only for the UK can an effect be perceived at all. Whereas for neither of the two countries have levels of short-run capital mobility been reached subsequently that are comparable to those that prevailed under the classical gold standard, long-run capital mobility seems to have been relatively high and - with the exception of the WWI-experience - also relatively constant over the whole century.

3.6 Conclusion

In this paper we have investigated in what sense correlations between savings and investment are informative about international capital mobility. Our reasoning uses insights from the theory of cointegrated systems and permanent-transitory decompositions and demonstrates that time series correlations between savings and investment are *per se* uninformative about the degree of international capital mobility. The findings of Feldstein and Horioka (1980) can therefore be rationalized even when capital mobility is perfect.

Even though this result is not new and has been put forward in the literature, the advantage of our approach is that we derive these conclusions from the reduced-form implications of an intertemporal maximization model. Hence, the results prevail independently of assumptions about the structure of underlying economic shocks. In particular, the implications of error correction for the cyclical dynamics of s and i have to our knowledge not been spelled out.

Still, the suggestion made by Feldstein and Horioka to make inference about international capital mobility from savings and investment data alone remains appealing. After

all, the theory *does* suggest that investments should flow where they yield the highest real returns and that savings depend on the intertemporal consumption decision alone.

In this paper, we have argued that the long-run adjustment process in a cointegrated system is informative about capital mobility . The adjustment coefficients also put us in a position to distinguish between (long-run) capital inflow and outflow mobility. We have also suggested a composite measure of long run capital mobility that arises naturally in the context of a cointegrated model and can be calculated easily as a by-product of Johansen's (1988) procedure. The measure has the advantage that it represents a standardized index of international capital mobility that is between zero and one. Also, standard errors of this index can be calculated and hence it becomes possible to compare capital mobility intertemporally and between countries.

Finally, we have applied our insights to a unique data set of historical savings and investment rates for the United States and the United Kingdom.The data are taken from Taylor (1996).

In the United States and the United Kingdom, long-run capital mobility over the century seems to have been remarkably high. WWI appears as the major disruption to long-run capital mobility in this century but in both countries long-run sustainable current account positions were restored soon after the war.

Whereas these findings seem somewhat at odds with the literature, we show that they are due to the fact that earlier studies tended to entangle the short and long-run dynamics of savings and investment. Our approach allows us to show that variations in capital mobility over the century have largely been reflected in changes in the short-run savings retention coefficient and whereas long-run capital mobility has been fairly high throughout the whole century.

This paper has concentrated on what we consider the essence of the Feldstein-Horioka approach: the claim that inference on capital mobility is possible from savings and investment data alone. We have demonstrated that this approach is valid if the appropriate reduced form that is suggested by the theory, i.e. a vector error correction model, is

74

chosen.

Bibliography

[1] Artis, M. J. and Bayoumi, T. (1992), 'Global Capital Market Integration and the Current Account,' in M.P. Taylor (Ed.) *Money and Financial Markets*, Cambridge, MS and Oxford: Blackwell 1992: 297-307.

[2] Bayoumi, T. (1990), 'Saving-Investment Correlations,' *IMF Staff Papers* vol. 37 no. 2: 360-87.

[3] Beveridge, S. and Nelson, C. R. (1981), 'A New Approach to the Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to the Measurement of the Business Cycle,' *Journal of Monetary Economics* 7: 151-74.

[4] Baxter, M. and Crucini, M. J. (1993), 'Explaining saving-investment correlations,' *American EconomicReview* 83: 416-436.

[5] Blanchard, O. and Quah, D. (1989), 'The Dynamic Effects of Demand and Supply Disturbances', *American Economic Review* 79: 655-73.

[6] Campbell, J. Y, and Shiller, R.J. (1987), ' Cointegration and tests of present value models,' *Journal of Political Economy* 95: 1062-1088.

[7] Coakley, J., Kulasi, F. and Smith, R. (1998), 'The Feldstein-Horioka Puzzle and Capital Mobility: A Review,' *International Journal of Finance and Economics* 3: 169-188.

[8] Dooley, M., Frankel, J. and Mathieson, (1987), 'International Capital Mobility: What Do Savings-Investment Correlations Tell us?', *Staff Papers of the International Monetary Fund*: 503-530.

[9] Eichengreen, B. (1990), 'Trends and Cycles in Foreign Lending', in H. Siebert (ed.), *Capital Flows in the World Economy*. Tuebingen: Mohr 1990.

[10] Eichengreen, B. (1992), *Golden Fetters*. Oxford: Oxford University Press.

[11] Engle, R. F., Hendry, D.F. and Richard, J.-F. (1983), 'Exogeneity', *Econometrica* 51: 277-304.

[12] Engle, R. and Granger, C. W. (1987), 'Co-Integration and Error Correction: Representation, Estimation and Testing' *Econometrica* 55: 251-76.

[13] Feldstein, M. and Horioka, C. (1980), 'Domestic Saving and International Capital Flows', *Economic Journal* 90: 314-29.

[14] Feldstein, M. and Bacchetta, P. (1991), 'National Savings and International Investment', in Bernheim, B. D. and Shoven, J. B. (eds.): *National Saving and Economic Performance*. Chicago: University of Chicago Press.

[15] Frankel, J. A. (1992), 'Measuring International Capital Mobility: A Review', *American Economic Review Papers and Proceedings*: 197-202.

[16] Ghosh, A. (1995), 'Capital mobility amongst the major industrialized countries: Too little or too much?,' *Economic Journal* 105: 107-28

[17] Glick, R. and Rogoff, K. (1995), 'Global versus country-specific productivity shocks and the current account,' *Journal of Monetary Economics* 35: 159-92.

[18] Gonzalo, J. and Granger, C. (1995), 'Estimation of Common Long-Memory Components in Cointegrated Systems,' *Journal of Business and Economics Statistics* 13: 27-35.

77

[19] Hansen, H. and Johansen, S. (1998),'Some Tests for Parameter Constancy in Cointegrated VAR-Models', European University Institute mimeo, July 1998.

[20] Johansen, S. (1988), 'Statistical Analysis of Cointegrating Vectors', *Journal of Economic Dynamics and Control* 12: 231-54.

[21] Johansen, S. (1991), 'Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vectorautoregressive models', *Econometrica* 59: 1551-80

[22] Johansen, S., (1997a), 'Common Trends in the Cointegrated VAR-model,' manuscript, European University Institute.

[23] Johansen, S. (1997b),'What is a positive relation?,' manuscript, European University Institute.

[24] Kasa, K. (1992), 'Common stochastic Trends in international stock markets,' *Journal of Monetary Economics* 29: 95-124.

[25] Mendoza, E.G. (1991), 'Real Business Cycles in a Small Open Economy,' *American Economic Review* 81: 797-818.

[26] Moreno, R. (1997), 'Saving-Investment Dynamics and Capital Mobility in the US and Japan', *Journal of International Money and Finance* 16: 837-863.

[27] Obstfeld, M. (1986), 'Capital Mobility in the World Economy: theory and measurement', *Carnegie Rochester Conference Series on Public Policy* vol 24: 55-104.

[28] Obstfeld, M. (1995), 'International Capital Mobility in the 1990s' in Kenen, P. (ed.) *Understanding Interdependence*, Princeton University Press, 1995.

[29] Obstfeld, M. and Rogoff, K. (1995), 'The intertemporal approach to the current account,' NBER working paper 4893.

[30] Proietti, T. (1997), 'Short-Run Dynamics in Cointegrated Systems,' *Oxford Bulletin of Economics and Statistics* 59: 405-422.

[31] Sachs, J. (1981), 'The Current Account and Macroeconomics adjustment in the 1970s,' *Brookings Papers on Economic Activity* 1: 201-268.

[32] Sarno, L. and Taylor M. P. (1998), 'Savings-Investment Correlations: transitory versus permanent,' *The Manchester School Supplement* 1998: 17-38.

[33] Taylor, A. M. (1994), 'Domestic Saving and International Capital Flows Reconsidered', NBER Discussion Paper Series, no. 4892.

[34] Taylor, A. M. (1996), 'International Capital Mobility in History: The Saving-Investment Relationship', NBER Discussion Paper Series, no. 5743.

3.7 Figures and Tables

Table 1

	Trace Statistics	Max EV Statistics
a)	cointegration tests for the US 1874-199	
$0 < h \leq 1$	22.29	16.73
$1 < h \leq 2$	5.564	5.564
b)	cointegration tests for the UK 1850-1992	
$0 < h \leq 1$	13.91	11.37
$1 < h \leq 2$	2.54	2.54
c)	UK 1850-92 with dummies for WWI&II	
$0 < h \leq 1$	59.3	56.96
$1 < h \leq 2$	2.34	2.35
Critical values 10% (5%)	$0 < h \leq 1$	$1 < h \leq 2$
trace test:	15.58 (17.48)	6.69 (8.803)
max-Eigenvalue-test:	12.78 (14.6)	6.69 (8.083)

Table 2: Estimated cointegrating vectors $\beta' = \begin{bmatrix} 1 & \beta_2 \end{bmatrix}$

	US (1874-1992)	UK (1850-1992)
β_2	-0.85	-0.65

Table 3: Tests of $H_0 : \beta' = \begin{bmatrix} 1 & -1 \end{bmatrix}$

	US (1874-1992)	UK (1850-1992)
LR	3.22	1.96
p-value	0.07	0.16

80

Table 4	Index of International Capital Mobility, $\Lambda = \alpha \Sigma_{00}^{-1} \alpha$	
	UK	US
1880-1913	0.37	0.23
	$\begin{pmatrix} 0.15 & 0.53 \end{pmatrix}$	$\begin{pmatrix} 0.01 & 0.41 \end{pmatrix}$
1919-39	0.28	0.51
	$\begin{pmatrix} 0.02 & 0.47 \end{pmatrix}$	$\begin{pmatrix} 0.15 & 0.72 \end{pmatrix}$
1946-71	0.51	0.91
	$\begin{pmatrix} 0.24 & 0.68 \end{pmatrix}$	$\begin{pmatrix} 0.76 & 0.95 \end{pmatrix}$
1972-92	0.29	0.21
	$\begin{pmatrix} 0.02 & 0.49 \end{pmatrix}$	$\begin{pmatrix} 0.01 & 0.28 \end{pmatrix}$

95% lower and upper confidence bounds after Hansen and Johansen (1998) in brackets

Table 5: Cointegration Tests 1950-92

	Trace Statistics	Max EV Statistics
a)	cointegration tests for the US 1950-92	
$0 < h \leq 1$	15.88	9.231
$1 < h \leq 2$	6.645	6.645
b)	cointegration tests for the UK 1950-1992	
$0 < h \leq 1$	16.24	15.24
$1 < h \leq 2$	0.99	0.99
Critical values 10% (5%)	$0 < h \leq 1$	$1 < h \leq 2$
trace test:	15.58 (17.48)	6.69 (8.083)
max-Eigenvalue-test:	12.78 (14.6)	6.69 (8.083)

Table 6: Estimated cointegrating vectors $\beta' = \begin{bmatrix} 1 & \beta_2 \end{bmatrix}$ and test of $\beta_2 = 1$

	US (1950-1992)	UK (1950-1992)
β_2	-1.32	-0.85
LR	0.6893	0.7359
p-value	0.4064	0.391

Table 7   Index of International Capital Mobility, $\Lambda = \alpha \Sigma_{00}^{-1} \alpha$

	UK	US
1950-1992	0.30	0.19
	$\begin{pmatrix} 0.24 & 0.37 \end{pmatrix}$	$\begin{pmatrix} 0.16 & 0.23 \end{pmatrix}$

95% lower and upper confidence bounds after Hansen and Johansen (1998) in brackets

83
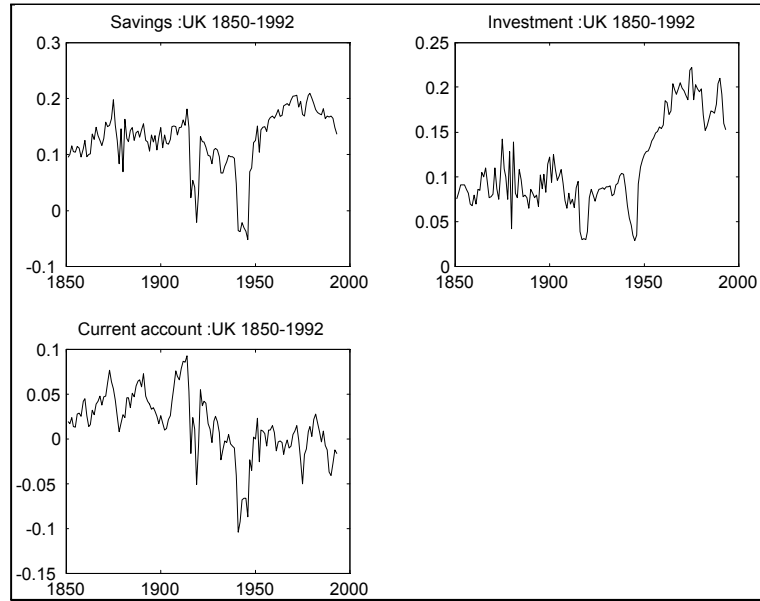
Figure 3-1: The UK Data 1850-1992

84
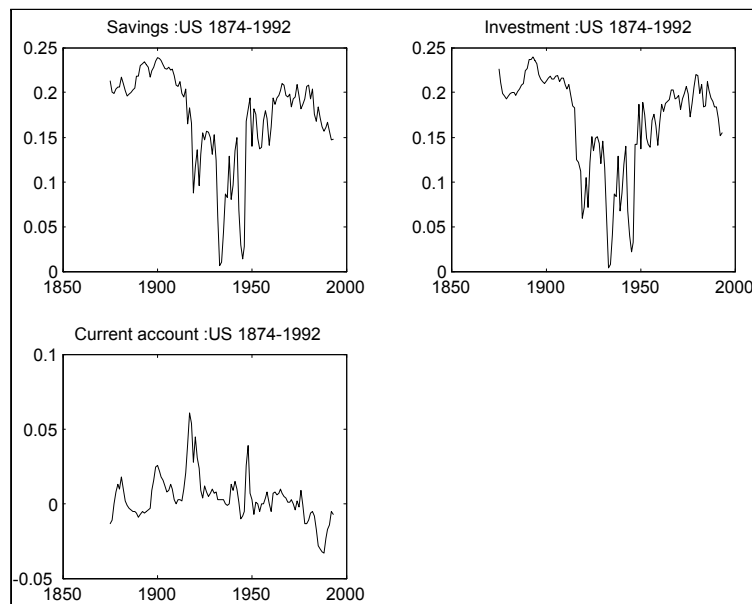
Figure 3-2: The US Data 1874-1992
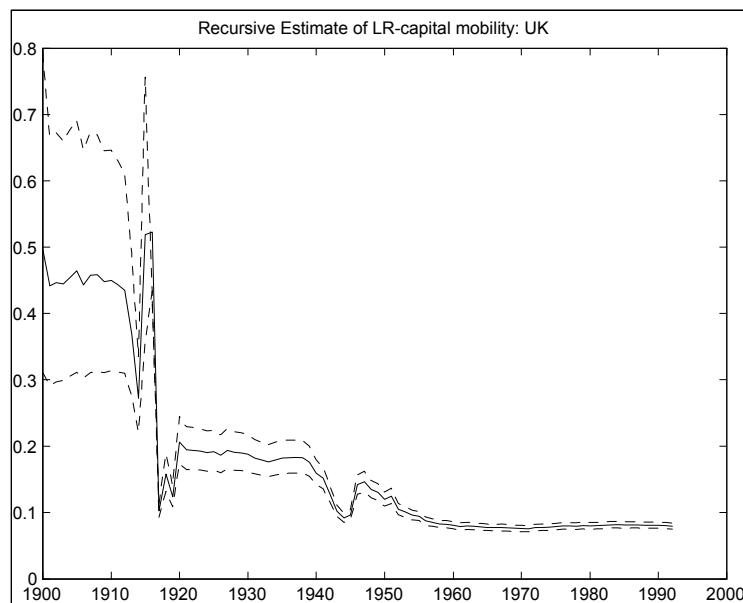
Figure 3-3: Long run capital Mobility in the UK

85
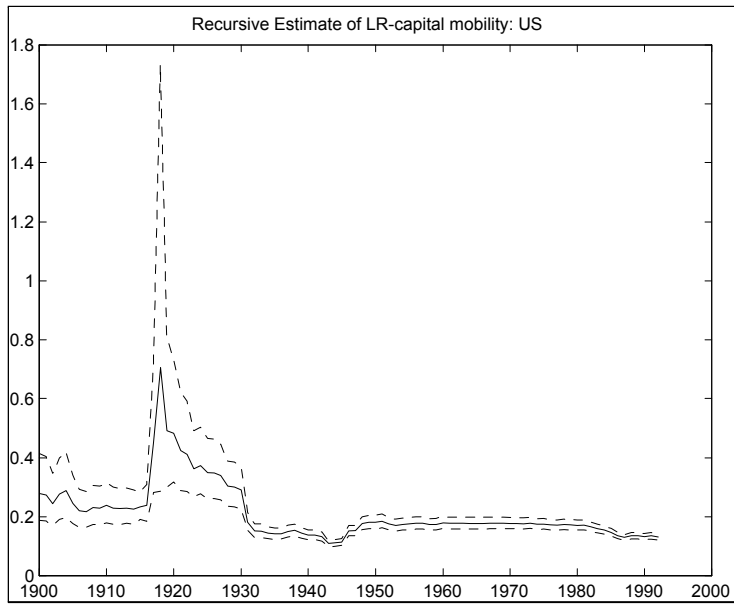
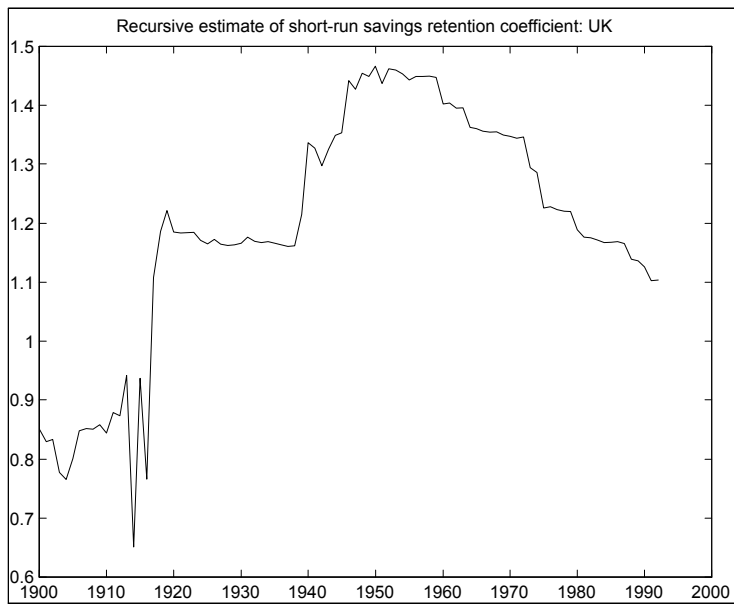Figure 3-4: Long run capital mobility in the U.S.
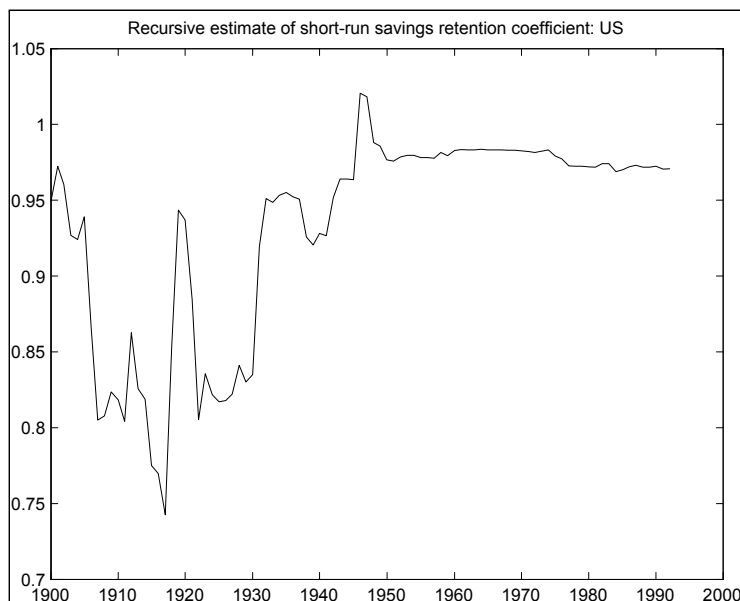

Figure 3-5: Short run capital mobility in the UK

86

Figure 3-6: Short run capital mobility in the U.S.

3.8 Mathematical Appendices

3.8.1 Appendix A

The permanent component according to Beveridge-Nelson (1981) is

$$\mathbf{X}_t^P = \mathbf{C}(\mathbf{1}) \sum_{l=1}^{t} \boldsymbol{\varepsilon}_l \tag{3.19}$$

where $\{\varepsilon_l\}$ is the series of innovations to $\mathbf{X}_t$ and $\mathbf{C}(\mathbf{1}) = \sum \mathbf{C}_i$ where the C_i are the coefficients of the moving-average (Wold) representation of $\Delta\mathbf{X}_t$. Now choose $\mathbf{X}_t = \begin{bmatrix} s_t & i_t \end{bmatrix}'$. It is important to recall that in the case where $\mathbf{X}_t$ has an error-correction representation, i.e. where

$$\Gamma(\mathbf{L})\Delta\mathbf{X}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{X}_{t-1} + \boldsymbol{\varepsilon}_t \tag{3.20}$$

87

there is a closed-form solution for the matrix $C(1)$, given by

$$\mathbf{C(1)} = \boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}_{\perp}'\Gamma(\mathbf{1})\boldsymbol{\beta}_{\perp})^{-1}\boldsymbol{\alpha}_{\perp}' \tag{3.21}$$

(See Johansen (1995)).

In our above model $\boldsymbol{\alpha} = \begin{bmatrix} \alpha_1 & \alpha_2 \end{bmatrix}'$ and $\boldsymbol{\beta} = \begin{bmatrix} 1 & -1 \end{bmatrix}$. Then $\boldsymbol{\alpha}_{\perp} = \begin{bmatrix} \alpha_2 & -\alpha_1 \end{bmatrix}'$ and $\boldsymbol{\beta}_{\perp} = \begin{bmatrix} 1 & 1 \end{bmatrix}$. Furthermore, let $\Gamma(1) = \begin{bmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & \gamma_{22} \end{bmatrix}$.

Plugging into the closed-form solution for $\mathbf{C(1)}$ yields

$$\begin{bmatrix} s_t \\ i_t \end{bmatrix}^P = A \begin{bmatrix} \alpha_2 & -\alpha_1 \\ \alpha_2 & -\alpha_1 \end{bmatrix} \sum_{l=0}^{t} \boldsymbol{\varepsilon}_l \tag{3.22}$$

where $A = 1/\left[(\gamma_{11} + \gamma_{12})\alpha_2 - (\gamma_{21} + \gamma_{22})\alpha_1\right]$ and hence

$$i_t^P = s_t^P$$

3.8.2 Appendix B

To derive our results, we draw heavily on work done by Johansen (1997), Proietti (1997) and Granger and Gonzalo (1995). We restate the VECM-representation:

$$\Gamma(\mathbf{L})\Delta\mathbf{X}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{X}_{t-1} + \boldsymbol{\varepsilon}_t \tag{3.23}$$

The transitory part of savings and investment is a moving average of reduced-form innovations (Beveridge-Nelson (1981)):

$$\begin{bmatrix} s_t - s_t^P \\ i_t - i_t^P \end{bmatrix} = \mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t$$

The idea is to approximate the transitory part by a linear combination of the current account. Premultiplying the VECM-representation by $\mathbf{C}(\mathbf{1})$ we obtain:

$$\mathbf{C}(\mathbf{1})\Gamma(\mathbf{L})\Delta\mathbf{X}_t = \mathbf{C}(\mathbf{1})\boldsymbol{\varepsilon}_t \qquad (3.24)$$

because $\mathbf{C}(\mathbf{1})\boldsymbol{\alpha} = \mathbf{0}$. Integrating yields:

$$\mathbf{C}(\mathbf{1})\Gamma(\mathbf{L})\mathbf{X}_t = \mathbf{C}(\mathbf{1})\sum_{l=0}^{t}\boldsymbol{\varepsilon}_l \qquad (3.25)$$

We now have a representation of the permanent component in terms of present and past levels of the process itself. Accordingly, we get for the transitory component:

$$\{I - C(1)\Gamma(L)\}\mathbf{X}_t = \mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t \qquad (3.26)$$

Let us now rewrite

$$\mathbf{C}(\mathbf{1})\Gamma(\mathbf{L}) = \mathbf{C}(\mathbf{1})\Gamma(\mathbf{1}) + \Delta\mathbf{C}(\mathbf{1})\Gamma^*(\mathbf{L})$$

where $\Gamma_i^* = -\sum_{j>i}\Gamma_j$. Then, in the above, we obtain:

$$\mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t = \{I - C(1)\Gamma(1)\}\mathbf{X}_t - \mathbf{C}(\mathbf{1})\Gamma^*(\mathbf{L})\Delta\mathbf{X}_t \qquad (3.27)$$

It is worthwhile to contemplate this result for a second. The transitory component

89

is a linear combination of the levels of the process plus some moving average of past changes. Note in particular, that $\{I - C(1)\Gamma(1)\}$ has rank $n - h = 1$ where $n = 2$ is the dimension of the system and $h = 1$ the number of cointegrating relations. Hence, the components of $\{I - C(1)\Gamma(1)\}\,\mathbf{X}_t$ are perfectly correlated, but the correlation can be both positive and negative. It is also important to note that $\{I - C(1)\Gamma(1)\}\,\mathbf{X}_t$ is just a linear combination of the equilibrium error $\boldsymbol{\beta}'\mathbf{X}_t = CA_t$. This can be seen from the following representation of the matrix $\{I - C(1)\Gamma(1)\}$ which has been derived by Proietti (1997):

$$\mathbf{I} - \mathbf{C(1)}\Gamma(\mathbf{1}) = (\Gamma(1) + \boldsymbol{\alpha\beta'})^{-1}\boldsymbol{\alpha}\left[\boldsymbol{\beta'}\left(\Gamma(1) + \boldsymbol{\alpha\beta'}\right)^{-1}\boldsymbol{\alpha}\right]^{-1}\boldsymbol{\beta'} = \boldsymbol{\psi\beta'} \qquad (3.28)$$

The expression $\{I - C(1)\Gamma(1)\}\,\mathbf{X}_t$ therefore captures the error correction mechanism of the model and we can rewrite:

$$\{\mathbf{I} - \mathbf{C(1)}\Gamma(1)\}\,\mathbf{X}_t = \boldsymbol{\psi\beta'}\mathbf{X}_t = \boldsymbol{\psi}CA_t \qquad (3.29)$$

For the second expression on the RHS of (3.27), we can write

$$\mathbf{C(1)}\Gamma^*(\mathbf{L})\Delta\mathbf{X}_t = \boldsymbol{\beta}_\perp f_t \text{ where } f_t = (\alpha'_\perp\Gamma(1)\beta_\perp)^{-1}\boldsymbol{\alpha}'_\perp\Gamma^*(\mathbf{L})\Delta\mathbf{X}_t$$

Here, f_t is a common factor and, since $\boldsymbol{\beta}_\perp = \begin{bmatrix} 1 & 1 \end{bmatrix}'$, the components of $\mathbf{C(1)}\Gamma^*(\mathbf{L})\Delta\mathbf{X}_t$ will be perfectly positively correlated.

Chapter 4

Current Accounts and the Persistence of Global and Country-Specific Shocks: Is Investment Really too Volatile?

4.1   Introduction

A better understanding of the empirical dynamics of the current account and investment in response to global and country-specific shocks is important as it puts to a test the modern 'intertemporal theory of the current account' (Obstfeld (1986, 1995), Sachs (1981), Obstfeld and Rogoff (1995)). Even though this theory is nowadays a theoretical workhorse in international macroeconomic analyses, empirical work in this area has so far been very sparse.

One exception is the important paper by Glick and Rogoff (1995). These authors empirically examined the role of global and country-specific productivity shocks for current account dynamics using a structural econometric model derived from the theory. Intertemporal optimization models predict that the current account reacts primarily to

91

country-specific shocks, not to global shocks: global shocks hit all economies equally and change consumption possibilities world-wide. Hence, there is no role for international borrowing and lending with a view to smoothing consumption. In response to a, say, negative country-specific shock, however, a country can borrow from the rest of the world in order to smooth consumption.

Overall, Glick and Rogoff could confirm these predictions of the theory. They found, however, that the reaction of investment to country-specific shocks was excessive vis-a-vis the implied current account response.

The puzzle encountered by Glick and Rogoff illustrates an important property of rational expectations models: their predictions crucially depend on whether structural shocks have permanent or transitory effects and also on the speed of adjustment to the new steady state (persistence). This sensitivity constitutes a dilemma for the empirical researcher: using univariate methods, it is almost impossible to distinguish between very persistent but stationary processes on one hand and unit-root processes on the other.

In this paper we suggest measuring the permanent component of shocks by choosing an appropriate VAR specification of the model and by exploiting cointegration in the data. In so doing, we can give a coherent description of the permanent and transitory components of global and country specific shocks with respect to the information set implied by the theory. Using this approach, we offer an alternative solution to the Glick and Rogoff puzzle: the current account seems to react stronger than investment to the permanent component of country-specific shocks but country-specific shocks have important transitory components. To the degree that these transitory components are not taken care of in the estimation of the impact response of savings and investment, estimates will be an amalgam of the response to transitory and permanent shocks. The kind of excess sensitivity of the current account response to varying degrees of persistence suggested by Glick and Rogoff as a solution to the puzzle is generally not empirically warranted. Our findings rather suggest an open-economy analogue of the solution proposed by Quah (1990) for the excess-smoothness of consumption: if economic agents distinguish between

92

transitory and permanent movements in their future income stream, low current-account investment correlations can be rationalized even if the current account is more sensitive to (the persistent component of) permanent shocks than is investment.

Our approach forces us to sacrifice some structure vis-a-vis the simultaneous equation model suggested by Glick and Rogoff. It is certainly a big advance of their study that the estimating equations are derived explicitly from an intertemporal model. The authors claim:

> 'The ability to derive closed-form solutions helps clarify some interesting issues that may easily be obscured in simulation analysis or vectorautoregression estimation' (Glick and Rogoff, pp.185-6)

In this study, we will argue that our understanding of current account and investment dynamics can be enhanced if both the economic theory as well as its reduced form are taken seriously. Employing a structural VAR, we use insights from the intertemporal model that are also confirmed by the results of Glick and Rogoff to identify country-specific and global shocks from the data directly. Using the same model framework, we also identify permanent and transitory shocks to investment and the current account. We are then able to describe the mapping between permanent and transitory shocks on the one hand and global and country-specific shocks on the other. Our reasoning will be based on geometric insights and will give rise to a measure of persistence of country-specific shocks. The quality of our identification of both country-specific shocks and their persistence is then assessed in two ways: first, cross-country-correlations of shocks are calculated for the panel of the seven largest economies in the world. Secondly, we use our models and the knowledge about country-specificity to forecast the current account based on a present value formula. Indeed, our models perform very well in forecasting current account behaviour.

The remainder of this paper is structured as follows: in section two we present the model of Glick and Rogoff (1995) and discuss how they derive the structural estimation

93

equations. In section 3 we will introduce our own approach. We suggest how to estimate permanent and transitory shocks as well as global and country-specific shocks from the data and we present a measure of persistence of country-specific shocks that is based on a geometric reasoning. Section 4 presents data and estimation results and section 5 concludes.

4.2 Structural estimation equations

Glick and Rogoff (1995) use a simple intertemporal model with adjustment costs and quadratic utility. The representative agent maximizes

$$E_t \sum_{i=0}^{\infty} \left(\frac{1}{r}\right)^i U(C_{t+i}) \text{ where } U(C) = C - \frac{h}{2}C^2 \tag{4.1}$$

subject to the intertemporal budget constraint

$$B_{t+1} = RB_t + NO_t - C_t \tag{4.2}$$

where $[B, NO, C]$ denote the net foreign asset position, net output defined as the difference between GDP and Investment, $NO_t = Y_t - I_t$, and consumption respectively and $R = 1 + r$ where r is the world interest rate which here is assumed to equal the representative individual's rate of time preference. The current account is then given by the change in the net foreign asset position, $CA_t = \Delta B_t$. Equivalently, defining saving as $S = Y - C + rB$ we get the conventional definition of the current account, $CA = S - I$.

The production side of the economy is described by a Cobb-Douglas type production function given by

$$Y_t = A_t^c A_t^w K_t^\gamma \left[1 - \frac{g}{2} \left(\frac{I_t^2}{K_t} \right) \right]$$

Here, K_t denotes the time t capital stock, $I_t = \Delta K_{t+1}$ is gross investment , γ is the

94

capital share of the economy, g is a positive constant and $\mathbf{A} = \begin{bmatrix} A_t^c, & A_t^w \end{bmatrix}'$ is a vector of country-specific and global total factor productivities which is supposed to follow an AR(1)-process:

$$\mathbf{A}_t = \begin{bmatrix} A_t^c \\ A_t^w \end{bmatrix} = \begin{bmatrix} \rho_{GR} & 0 \\ 0 & 1 \end{bmatrix} \mathbf{A}_{t-1} + \begin{bmatrix} \varepsilon_t^c \\ \varepsilon_t^w \end{bmatrix} \tag{4.3}$$

where ε_t^c and ε_t^w are supposed to be mutually uncorrelated at all leads and lags.

Glick and Rogoff (1995) linearize the first order conditions which yields a system of equations of the following form:

$$Y_t = a_I I_t + a_K K_t + \mathbf{a}_A' \mathbf{A}_t + \mu_{Yt} \tag{4.4}$$

$$I_t = b_1 I_{t-1} + \sum_{s=1}^{\infty} \boldsymbol{\lambda}_s' \{E_t \mathbf{A}_{t+s} - E_{t-1} \mathbf{A}_{t+s-1}\} + \mu_{It} \tag{4.5}$$

$$C_t = \frac{R-1}{R} \left(B_t + E_t \sum_{s=0}^{\infty} R^{-s} NO_{t+s} \right) + \mu_{Ct} \tag{4.6}$$

where $\boldsymbol{\lambda}_s' = \begin{bmatrix} d_c \lambda_c^s, & d_w \lambda_w^s \end{bmatrix}$ and ,λ_c and λ_w are positive and smaller than unity.

In the above, $\boldsymbol{\mu}' = \begin{bmatrix} \mu_{Yt} & \mu_{It} & \mu_{Ct} \end{bmatrix}$ is a vector of mutually uncorrelated i.i.d. disturbances that is added *ad hoc* to provide the error structure for the estimation equations. From this linearization, it is then possible to derive the estimable equations

$$\Delta I_t = (b_1 - 1) I_{t-1} + b_2 \Delta A_t^c + b_3 \Delta A_t^w + v_{It} \tag{4.7}$$

95

and in the case of $\rho_{GR} = 1$:

$$\Delta CA_t = c_1 I_{t-1} + c_2 \Delta A_t^c + c_3 \Delta A_t^w + r CA_{t-1} + v_{CAt} \tag{4.8}$$

Again, $\mathbf{v}_t' = \begin{bmatrix} v_{It} & v_{CAt} \end{bmatrix}$ are error terms that are functions of $\boldsymbol{\mu}_t$. Glick and Rogoff also show that v_{CAt} is correlated with CA_{t-1} whereas I_{t-1} is predetermined in the equation for ΔCA_t. They solve this problem by imposing a value for r. Then the system of equations (4.7) and (4.8) can be estimated by two stage least squares as a seemingly unrelated regression model.

It is an important result of Glick and Rogoff that the coefficient on ΔA_t^w in the CA-equation is found to be insignificant for all seven countries, in accordance with the theory. However, their empirical implementation reveals a puzzle:

Under the assumption that country-specific shocks do have a permanent effect on net output, the theory also predicts that $|c_2|/b_2 > 1$, i.e. the reaction of the current account to country-specific shocks should be stronger than the response of investment. A positive, permanent country-specific TFP-shock increases today's gross output, Y_t. Future gross output will however even be higher than today's gross output because the productivity shock makes it profitable to invest. Hence the future capital stock and consequently also future output will be higher. Because consumption instantaneously adjusts to the permanently higher future output stream, this implies that savings will have to fall and hence the current account should change by more than investment (in the opposite direction, though).

From the data, Glick and Rogoff consistently find estimates of c_2 that are smaller in absolute value than those for b_2. This is puzzling but this result strongly depends on the persistence of country-specific shocks. Glick and Rogoff show that even for small deviations of ρ_{GR} from unity, the relative current-account / investment response can be substantially muted: as the shock is no longer permanent, people will save more instead of less. At the same time, the incentive to invest is weakened as productivity will only be temporarily high. Glick and Rogoff show that for reasonably chosen parameter values

96

of the structural model the CA/I response will fall into the range of their estimates.

In the following section, we outline an alternative approach that relies on measuring the relative importance of transitory and permanent components in country-specific shocks rather than specifying it *a priori*, as in equation (4.3) which requires shocks to be fully permanent or fully transitory.. As we will show, our more data-driven approach leads to an alternative solution of the Glick-Rogoff puzzle: if shocks have both permanent and transitory components, the estimated response in the Glick-Rogoff model may be an amalgam of responses to permanent and transitory shocks.

Our method is based on a cointegrated VAR-model of investment and the current account: we first identify global and country-specific shocks from the data. Then we rerun the model with an alternative identification scheme that exploits the cointegrating information in the data to identify permanent and transitory shocks. We are then able to compare global and country-specific shocks with permanent and transitory disturbances and we can investigate how one class of shocks maps into the other. We can then suggest a measure of the persistence of global and country-specific shocks that is based on a geometric reasoning.

4.3   Identifying the shock matrix

In this section, we will present the structural VAR techniques that we will use to measure the persistence of country -specific and global shocks.

We will consider a simple bivariate VAR in investment and the current account:

$$\Pi(\mathbf{L})\mathbf{X}_t = \varepsilon_t \tag{4.9}$$

where $\mathbf{X}_t' = \begin{bmatrix} CA_t, & I_t \end{bmatrix}$.

Our aim is to identify two classes of shocks from this model: permanent vs. transitory shocks and global vs. country-specific shocks. Furthermore, we want to find out how one class of shocks maps into the other, i.e. we want to know how persistent country-specific

97

shocks are or we want to know how much of the typical variation in permanent shocks is explained by global influences.

4.3.1   Permanent vs transitory

If investment and savings in the model-economy laid out in section 2 can be characterized by $I(1)$-processes, then the intertemporal approach imposes a cointegrating relationship on the data: the current account will have to be stationary as it can be represented as the discounted sum of changes in net output. As net output is itself assumed to be an $I(1)$-process, its differences will be $I(0)$ and so will be the current account. As investment and savings are $I(1)$, there is a cointegrating relationship between them.

Cointegration is a general property of present value models and the implications of this property for econometric modelling have first been explored by Campbell and Shiller (1987). In our model, the cointegrating restriction amounts to saying that CA_t is stationary while I_t is not. Let us rewrite the VAR in error correction form (VECM), neglecting constant terms:

$$\Gamma(\mathbf{L})\Delta\mathbf{X}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{X}_{t-1} + \boldsymbol{\varepsilon}_t \tag{4.10}$$

Then the theory would predict that $\boldsymbol{\beta}' = \begin{bmatrix} 1 & 0 \end{bmatrix}$.

The VECM can be inverted to yield a Beveridge-Nelson-Stock-Watson (BNSW) representation in terms of reduced-form disturbances:

$$\mathbf{X}_t = \mathbf{C(1)} \sum_{l=0}^{t} \boldsymbol{\varepsilon}_l + \mathbf{C^*(L)}\boldsymbol{\varepsilon}_t \tag{4.11}$$

where $C^*(L)\varepsilon_t$ is a stationary moving average and the first term is the random walk component of the $I(1)$-process X_t. As Johansen (1995) has shown, $C(1)$ has a closed-form

98

representation in terms of the parameters of the VECM:

$$\mathbf{C(1)} = \boldsymbol{\beta}_\perp (\boldsymbol{\alpha}'_\perp \Gamma(\mathbf{1}) \boldsymbol{\beta}_\perp)^{-1} \boldsymbol{\alpha}'_\perp \tag{4.12}$$

where $\boldsymbol{\alpha}_\perp, \boldsymbol{\beta}_\perp$ are the orthogonal complements of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ respectively. As this representation shows, $\mathbf{C(1)}$ is of reduced rank: if there are h cointegrating relationships, then $\mathbf{C(1)}$ has rank $n - h$ where n is the dimension of the system. This reflects the fact that in a cointegrated system, there is a reduced number of common trends that drive the system in the long-run. This is what underlies the Stock-Watson representation of a cointegrated stochastic process. We can write the random walk component as

$$\mathbf{C(1)} \sum_{l=0}^{t} \boldsymbol{\varepsilon}_l = \mathbf{A}_0 \boldsymbol{\alpha}'_\perp \sum_{l=0}^{t} \boldsymbol{\varepsilon}_l = \mathbf{A}_0 \boldsymbol{\tau}_t \tag{4.13}$$

where the common trends are given by $\boldsymbol{\tau}_t = \boldsymbol{\alpha}'_\perp \sum_{l=0}^{t} \boldsymbol{\varepsilon}_l$. Accordingly, the permanent shocks to the system are just given by $\boldsymbol{\eta}_t = \boldsymbol{\alpha}'_\perp \boldsymbol{\varepsilon}_t$. If we require that permanent and transitory shocks should be orthogonal to each other, the transitory shocks are given by

$$\xi_t = \boldsymbol{\alpha}' \Omega^{-1} \boldsymbol{\varepsilon}_t \tag{4.14}$$

where Ω is the variance-covariance matrix of the reduced form residuals $\boldsymbol{\varepsilon}_t$.

Hence, the matrix $\mathbf{P}$ that maps $\boldsymbol{\varepsilon}_t$ on the vector of permanent and transitory disturbances, $\boldsymbol{\theta}'_t = \begin{bmatrix} \eta_t, & \xi_t \end{bmatrix}$ is given by

$$\mathbf{P} = \begin{bmatrix} (\alpha'_\perp \Omega \alpha_\perp)^{-1/2} \, \boldsymbol{\alpha}'_\perp \\ (\alpha' \Omega^{-1} \alpha)^{-1/2} \, \boldsymbol{\alpha}' \Omega^{-1} \end{bmatrix} \tag{4.15}$$

where the factors $(\alpha'_\perp \Omega \alpha_\perp)^{-1/2}$ and $(\alpha' \Omega^{-1} \alpha)^{-1/2}$ normalize η_t and ξ_t to have unit variance.

99

4.3.2 Global vs. country-specific

We are now in a position to identify permanent and transitory disturbances. In a next step, we need to identify global and country-specific shocks from the data. The solution in this case will not come from a correct interpretation of the parameters of the econometric model but rather from outside, i.e. from economic theory. Theory predicts that the current account should not react to global shocks. Our tests will be flawed if we wrongly build our analysis on this presumption. But this is exactly the main finding by Glick and Rogoff: in their estimates, the global shock almost never has a significant effect on the current account in the same period. We can therefore base our analysis on theirs, assuming that we can validly identify global from country-specific shocks by imposing that the former do not have a contemporaneous effect on the current account.

In the framework of our VAR, this amounts to a very simple and convenient identifying restriction: identification is achieved by means of a Choleski decomposition of the variance-covariance matrix of the reduced form residuals, Ω. To see this, consider the BNSW-representation

$$\mathbf{X}_t = \mathbf{C}(\mathbf{1}) \sum_{l=1}^{t} \boldsymbol{\varepsilon}_l + \mathbf{C}^*(\mathbf{L})\boldsymbol{\varepsilon}_t \tag{4.16}$$

we have referred to this as the 'reduced' form. We can rewrite in difference form:

$$\Delta\mathbf{X}_t = \mathbf{C}(\mathbf{L})\boldsymbol{\varepsilon}_t \tag{4.17}$$

where the coefficients of the matrix polynomial $\mathbf{C}(\mathbf{L})$ are given by $\mathbf{C}_i = \mathbf{C}_i^* - \mathbf{C}_{i-1}^*$. Then, we hypothesize the existence of a structural form

$$\Delta\mathbf{X}_t = \mathbf{D}(\mathbf{L})\mathbf{e}_t$$

where $\mathbf{e}_t' = \begin{bmatrix} e_t^c, & e_t^w \end{bmatrix}$ is the vector of country-specific and global shocks. It is assumed

100

that the reduced form residuals are a linear function of the structural disturbances e_t:

$$\varepsilon_t = \mathbf{S}e_t \tag{4.18}$$

Furthermore, the structural disturbances are orthonormal, i.e. $\mathbf{var}(\mathbf{e}_t) = \mathbf{I}_n$. It is then clear that

$$\Omega = \mathbf{SS}' \tag{4.19}$$

and

$$\mathbf{D(L)} = \mathbf{C(L)S} \tag{4.20}$$

In a bivariate system, the first of these conditions gives three restrictions for the four elements of $\mathbf{S}$. To achieve identification, one additional restriction is needed and we get it from the theory: global shocks do not have a contemporaneous impact on the current account. Recalling that $\mathbf{X}_t' = \begin{bmatrix} CA_t & I_t \end{bmatrix}$ and

$$\mathbf{C(0)S} = \mathbf{S} = \mathbf{D(0)} \tag{4.21}$$

this amounts to assuming that S is lower triangular:

$$\mathbf{S} = \begin{bmatrix} s_{11} & 0 \\ s_{12} & s_{22} \end{bmatrix}$$

We now have classified disturbances to our bivariate system according to two categories: their persistence and their country-specificity. The question that we set out to answer is: how persistent are country-specific and global shocks? We are now in the position to answer this question. The matrix that maps global and country-specific shocks

101

into the permanent and transitory domain is given by

$$\boldsymbol{\theta}_t = \mathbf{PS}\mathbf{e}_t = \mathbf{Q}\mathbf{e}_t \tag{4.22}$$

Note that $\mathbf{Q} = \mathbf{PS}$ is a orthonormal matrix, i.e. $\mathbf{QQ}' = \mathbf{I}_n$.

The matrix $\mathbf{Q}$ contains all the information we are interested in. In fact, $\mathbf{Q}$ is nothing else than the covariance of $\boldsymbol{\theta}_t$ and $\mathbf{e}_t$:

$$E(\boldsymbol{\theta}_t \mathbf{e}_t') = \mathbf{Q}E(\mathbf{e}_t \mathbf{e}_t') = \mathbf{Q} \tag{4.23}$$

Note that due to the unit variance of the components of $\mathbf{e}_t$ and $\boldsymbol{\theta}_t$, $\mathbf{Q}$ also defines the cross-correlation of $\mathbf{e}_t$ and $\boldsymbol{\theta}_t$ But beyond being covariance and correlation matrix at the same time, the orthonormality of $\mathbf{Q}$ provides a particular structure. It tells us, that if we choose the orthogonal basis of permanent and transitory shocks as our coordinate system, global and country-specific shocks are just a pair of orthogonal vectors in this coordinate system and the coordinates are given by the rows of $\mathbf{Q}$. Also, the squares of this coordinates are just the share of the variance of $\mathbf{e}_t$ that is given by permanent and transitory shocks. Figure (1) in the appendix, illustrates this geometric intuition: the upper left entry of $\mathbf{Q}$ which we will henceforth denote by ρ, is nothing else than the cosine of the angle λ between the typical country-specific shock and the permanent axis, the span of $[0, \eta]'$.

In fact, $\mathbf{Q}$ is nothing else than a rotation of the orthogonal basis of the country-specific and global shocks onto the basis of permanent and transitory shocks. Hence, the parameter ρ or, alternatively, the angle λ uniquely determine $\mathbf{Q}$. In other words: the space of orthonormal (2×2) matrices is one-dimensional. This becomes immediately apparent from recalling that $\mathbf{QQ}' = \mathbf{I}$, which imposes 3 non-redundant restrictions on $\mathbf{Q}$. We can then parametrize $\mathbf{Q}$ as a function of the permanent component of country-specific

102

shocks as follows:

$$\mathbf{Q}\left(\rho\right) = \begin{bmatrix} \rho & -\sqrt{1-\rho^2} \\ \sqrt{1-\rho^2} & \rho \end{bmatrix} = \begin{bmatrix} \cos\lambda & -\sin\lambda \\ \sin\lambda & \cos\lambda \end{bmatrix} \tag{4.24}$$

We deliberately choose ρ to denote the permanent components of country-specific shocks, in analogy to ρ_{GR} in section 2. Certainly, these are not the same parameters but in the context of different models they formalize the same notion: ρ measures the correlation between the country-specific and the permanent shock in the VAR, whereas ρ_{GR} roughly measures the conditional correlation between A_t^c and A_{t-1}^c. In this sense, both ρ and ρ_{GR} are persistence measures.

4.3.3   Current account response and persistence

Glick and Rogoff (1995) show theoretically, how the period zero current account and investment responses depend on the persistence of country-specific shocks. In this sub-section, we will discuss how our framework can be used to assess whether excess sensitivity can account for their results. Recall that Glick and Rogoff found that, empirically, investment reacts much stronger than the current account in response to a country-specific shock. In terms of our model, that corresponds to estimates of the matrix $\mathbf{S} = \{s_{ij}\}$ such that $|s_{11}| < s_{12}$. However, as long as country-specific shocks have some permanent impact, the prediction of the theory is just the inverse: the current account should react much stronger than investment.

Note that our measure of persistence, ρ and hence the matrix $\mathbf{Q}$ is a function of the period zero impulse response of current account and investment. Taking an 'inverse engineering' approach, we can therefore ask a question that is the reduced-form analogue to Glick and Rogoff: how does persistence depend on changes in the relative impulse responses and vice versa?

103

For this purpose, recall that

$$\mathbf{Q} = \mathbf{PS} \tag{4.25}$$

Now let $\boldsymbol{\alpha}' = \begin{bmatrix} \alpha_1 & \alpha_2 \end{bmatrix}$. Then $\boldsymbol{\alpha}'_\perp = \begin{bmatrix} -\alpha_2 & \alpha_1 \end{bmatrix}$. Furthermore, let $\Omega = \{\omega_{ij}\}$. Note also that $\mathbf{S}$ is just the lower Choleski-factor of Ω which is given by

$$\mathbf{S} = \begin{bmatrix} \sqrt{\omega_{11}} & 0 \\ \omega_{21}/\sqrt{\omega_{11}} & \sqrt{\omega_{22} - \omega_{21}^2/\omega_{11}} \end{bmatrix} \tag{4.26}$$

Then plugging in for $\mathbf{Q}$ we can write the upper left entry, ρ, as follows:

$$\rho = \frac{-\alpha_2\sqrt{\omega_{11}} + \alpha_1\omega_{21}/\sqrt{\omega_{11}}}{\sqrt{\alpha_2^2\omega_{11} + a_1^2\omega_{22} - 2\alpha_1\alpha_2\omega_{21}}} \tag{4.27}$$

Let us also consider the relative impulse response of current account and investment which from the above is just given by the ratio of the current account variance to the covariance with investment:

$$\chi = \frac{s_{11}}{s_{21}} = \frac{\omega_{11}}{\omega_{21}}$$

We now have expressed the persistence of country-specific shocks as an involved function of the adjustment coefficients $\boldsymbol{\alpha}$, the variance-covariance-structure of investment and the current account. However, we should rather think of ρ as the natural parameter and the impulse response and hence the covariance structure as an outcome of the economic structure. What we are particularly interested in is the change of the impulse response with respect to a change in persistence around $\rho = 1$.

The strategy we are going to pursue is as follows: we are going to reparameterize ρ and χ as functions of the correlation of current account and investment which is defined

104

by

$$\phi = \frac{\omega_{21}}{\sqrt{\omega_{11}\omega_{22}}}$$

I.e., we are going to treat the adjustment parameters, $\boldsymbol{\alpha}$, and the conditional variances of investment and the current account, ω_{22} and ω_{11} respectively, as fixed. The correlation ϕ therefore contains the same information as ω_{21}. Using the implicit function theorem, we can then express $\partial\phi/\partial\rho$ at $\rho = 1$ and therefore also get a notion of the sensitivity of χ in a neighbourhood of $\rho = 1$. This is done in the mathematical appendix. Before we provide the results, however, let us briefly sharpen our intuition by considering what happens if $\rho = 1$. We can then solve (4.27) to find that

$$\phi = \pm 1 \tag{4.28}$$

This is an important first result: if and only if country-specific shocks are completely persistent, we should expect changes in the current account and investment to be perfectly correlated. This explains why Glick and Rogoff - like many other authors - find a robust negative correlation that is, however, significantly different from one. Complete persistence of country-specific shocks leads to singularity of the matrix Ω, which is another way of stating that investment and the current account have a 'common cycle'[1].

In the appendix, we derive the following expression for $\partial\phi/\partial\rho$ at $\rho = 1$:

$$\partial\phi/\partial\rho_{|\rho=1} = \left[1 - \frac{\alpha_2}{\alpha_1}\sqrt{\frac{\omega_{11}}{\omega_{22}}}\right]^2 \tag{4.29}$$

---

[1]This is just a dual way of phrasing the Feldstein-Horioka puzzle: if changes in the current account actually represent changes in investment then the covariance between savings and investment changes will be zero. In earlier work (Hoffmann (1998)), we have argued that correlations of appropriately detrended savings and investment data can take any value without assumptions on the structure of underlying shocks but that they are *per se* uninformative about capital mobility. The present paper can be interpreted as extending this argument to *changes* of savings and investment: if country-specific shocks are not permanent but persistent, investment-current account relations can be low without any implications for capital mobility. In theoretical terms, this insight has first been put forward by Obstfeld (1986, 1995).

105

Plugging into $\chi = \frac{\sqrt{\omega_{11}}}{\phi\sqrt{\omega_{22}}}$ and doing a Taylor expansion around $\phi = 1$, we find that

$$\chi(1 - \Delta\rho) = \sqrt{\frac{\omega_{11}}{\omega_{22}}} + \sqrt{\frac{\omega_{11}}{\omega_{22}}} \left[1 - \frac{\alpha_2}{\alpha_1} \sqrt{\frac{\omega_{11}}{\omega_{22}}} \right]^2 \Delta\rho \qquad (4.30)$$

and obviously, we can approximate

$$\partial\chi/\partial\rho = \sqrt{\frac{\omega_{11}}{\omega_{22}}} \left[1 - \frac{\alpha_2}{\alpha_1} \sqrt{\frac{\omega_{11}}{\omega_{22}}} \right]^2$$

This is the second important result of this section: using the parameters of the reduced form, we can estimate, how sensitive the current-account and investment response would be to small changes in the persistence of country-specific shocks around $\rho = 1$ - keeping $(\boldsymbol{\alpha}, \Omega)$ fixed. This puts us in the position to empirically assess whether small departures from the assumption that country-specific TFP follows a random-walk can rationalize the findings of Glick and Rogoff.

4.3.4 Forecast performance and country-specificity

The essential message of the previous section was that small changes in persistence can have dramatic effects on the dynamic responses of investment and the current account.

In this section, we will argue that the forecast performance of VARs can be used to assess the validity of the intertemporal approach. This idea is not new. There is a developing but still small literature that tests the present value formula of the current account that is implied by the intertemporal approach (Sheffrin and Woo (1990), Gosh (1995)). The general flavour of the results is that VAR-forecasts based on a present value formula do a good job in tracking ups and downs in the current account (i.e. are highly correlated with observed current accounts). Yet. the volatility of the implied current account forecasts often differs markedly from the actually observed current account (for an illustration, see also the graphs in Obstfeld and Rogoff (1996), pp.93-95).

Let us now illustrate the procedure that is generally employed for current-account

forecasts: first, a bivariate VAR is estimated, consisting of the real current account and a proxy of net output, NO_t:

$$\mathbf{B(L)Z}_t = \boldsymbol{\varepsilon}_t \text{ where } \mathbf{Z}_t = \left[\begin{array}{cc} \Delta NO_t, & CA_t \end{array} \right]'$$

Then, the VAR is used to forecast ΔNO_t. The current account, in a simple model with quadratic utility like the one laid out above, can be expressed as the present discounted value of expected changes in net output:

$$CA_t = -\sum_{l=1}^{\infty} R^{-l} E_t(\Delta NO_{t+l}) \tag{4.31}$$

The VAR forecasts of ΔNO_{t+l} can be used to approximate agent's expectations and an implied current account can be calculated from the VAR, once a plausible value for the interest rate is imposed.

In some cases this procedure works well, while in others, it does a very bad job. The theory, however, makes much stronger statements about which changes in net output should drive the current account: it tells us that if capital markets are sufficiently integrated, then global shocks should not impinge on the current account at all. Based on our reasoning in the previous section, it may be possible to improve forecasts of the current account by taking into consideration only those predictable changes in net-output that are driven by country-specific shocks. Hence, we can restrict our forecast of changes in net output to the component that is driven by country-specific shocks. If we have identified country-specific shocks well and if our theory is compatible with the data, we should be able to forecast the current account at least as well as if we chose the traditional approach.

Even though we have considered a model that contains investment instead of net output, we are going to use investment as proxy of net output: if agents expect higher net output, they will invest more and hence changes in investment should be highly correlated with changes in net output. As we will see, this notion is also empirically justified and

107

in some cases, we are able to substantially improve over the naive (traditional) way of forecasting the current account.

4.4 Empirical results

4.4.1 Data and model specification

In the estimation of our model, we used the data given in the appendix of Taylor (1996): annual savings and investment rates for the G7-countries (Unites States, Japan, Germany, France, Italy, the United Kingdom and Canada) from 1960 to 1991. We then used the real GDP data in Gordon (1993) to convert the rates into levels.

In a first step, we estimated an unrestricted VAR in levels to determine the correct lag length of the VAR model. Hannan-Quinn-, Schwarz- and Akaike information criteria all suggested that one to two lags yielded an adequate representation for all of the countries. To allow for richer dynamics, we chose two lags for all models. We then performed tests for cointegration based on Johansen's (1988) procedure. In three cases we did not find cointegration: for the US, Canada and the UK, no cointegration could be detected, whereas for Japan cointegration was detected at the 90-percent significance level in the maximum eigenvalue test. However, our sample is quite short (31 observations) and the low power of unit-root tests in particular in small samples, is well known. Also, we have strong theoretical priors: a nation's intertemporal budget constraint will restrict its current account dynamics in the long run. We therefore decided to impose one cointegrating restriction in the estimation of all seven models.

For the United States, Germany and Japan we had difficulties in establishing that the current account is indeed stationary, rather, it seems that for those countries we have a non-trivial cointegrating relationship. However, it is difficult to conceive of a theoretically meaningful cointegrating relationship between the current account and investment. Rather, these results seem to suggest that there is an important variable missing. Figure 2 plots the cointegrating residuals for these three countries vis-a-vis the long-term

108

interest rate differential with the United States. Upon visual inspection, the correlation is striking and it seems to suggest that the dynamics of the current account for these countries cannot be adequately modelled without taking account of the common factor represented by the interest rate differential.

For Japan, the US and Germany, we therefore set up a trivariate VAR with the interest rate differential vis-a-vis the US (vis-a-vis Germany for the US). We detected one cointegrating relationship in all three cases. We then tested for weak exogeneity of the interest rate differential. This also was accepted in all three cases. We can therefore return to our bivariate VAR of current account and investment as a conditional model, treating the interest rate differential as an exogenous variable. Indeed, now the hypothesis that $\boldsymbol{\beta}' = [1,0]$ was accepted for both Japan and the United States. For Germany, the hypothesis still could not be accepted but a cointegrating vector of $[1, 1/2]$ seemed compatible with the data and we decided to model the German economy with this cointegrating vector imposed.

Also for Canada and the UK we decided to introduce conditioning variables: the oil price in the UK model and the Can\$/US\$ nominal exchange rate for the Candian model. This was done for reasons of forecast performance which will be discussed later in this section.

In tables 4.1 and 4.2 we report test results on our final model specifications, i.e. with the exogenous regressors included. For Japan and Germany, we now find cointegration at high significance levels and also the theoretical value of $\beta' = [1,0]$ is not rejected in tests on the cointegrating space, except in the German case.

Recent work by Harbo et al. (1998) has established that the distributions of tests for cointegrating rank in partial systems can be substantially altered vis-a-vis the standard distributions that arise when the partial system is treated as if it was a full system. Hence, our systems should be regarded as two-dimensional subsystems of three-dimensional systems where one variable does not react to the equilibrium error. Using the results from table 3 in Harbo et al. (1998) in our table 4.1 we now also accept cointegration for both

the UK and Canada.

4.4.2 Persistence and country-specificity

In Table 4.3 we give the estimates of the matrix **Q** for all countries. Note that there is nothing to prevent empirical estimates of ρ from becoming negative. The sign of ρ is without importance in our context, however and that is why we report values of ρ^2. This gives us the added benefit that ρ^2 can be interpreted as the share of permanent shocks in the variability of the country-specific shocks.

On average, global shocks seem to be primarily permanent whereas country-specific shocks are not very persistent. There are however, a few exceptions: For Japan, 38 percent of the variability in the country-specific shock seems to be explained by permanent influences. For Germany, the country-specific shock seems highly persistent as well, 86 percent of its variance are explained by permanent influences.

One clear result stands out, however: country-specific shocks are neither fully permanent nor completely transitory. On average, 23 percent of the variance of country-specific shocks is explained by permanent influences. Theoretical models, in which country-specific TFP follows either a random walk or is just a mean-reverting process are therefore likely to give misleading results.

We showed earlier that the persistence of country-specific shocks is also going to influence the immediate response of investment and the current account. In table 4.4 we give our estimates of the Choleski-factor **S** of the reduced form covariance matrix Ω. The result is striking; by and large, the Glick-Rogoff puzzle disappears: for most countries, the current account response is $1 - 2$ times stronger than the investment response. Also, in all cases, their signs are opposite. There, are two exceptions: the United States, where the puzzle persists and investment still reacts twice as strong as the current account and Italy where the ratio is slightly smaller than unity. For the UK, it is roughly equal to one. The results all share a common feature of SVAR impulse responses: the standard errors are very large. Nonetheless, it is an encouraging result that the point estimates are in the

110

range predicted by the theory. Also, calculating the average of the ratio s_{11}/s_{21} across all countries is we get a value of -1.23, clearly in the range predicted by the theory.

On the other hand, we can also take a counterfactual look at the implied response if country-specific shocks were completely permanent. This is given by $\chi(1) = -\omega_{11}/\omega_{22}$. Table 4.5 compares the Glick and Rogoff responses with the responses implied by our model at $\rho = 1$. Conversely, it also provides the implied persistence of the Glick and Rogoff response in terms of our model, which is given by $1 - \Delta\rho$, where $\Delta\rho$ can be calculated from the Taylor-approximation in (4.30). At first sight it seems that small departures from the random-walk assumption can account for the impulse responses found by Glick and Rogoff: on average our estimate of the implied ρ equals 0.95, very close to the 0.97 average autocorrelation coefficient in the original study. However, the estimated sensitivities are generally fairly low, so even though $\chi(1)$ is generally bigger in absolute value than the GR-estimate, assuming $\rho = 1$ only goes a small way towards bringing the impulse response into the range predicted by the theory. The average $\chi(1)$ - not including Canada - is -0.63. For Germany we find a rather high sensitivity and here the Glick-Rogoff approach goes furthest towards explaining the puzzle. Also for France, half of the difference between the GR-impulse response and unity can be bridged by letting ρ go to unity. For Canada we find a sensitivity close to zero which suggests that $\alpha_2/\alpha_1 \approx \sqrt{\omega_{22}/\omega_{11}}$, an unusual parameter constellation for which we do not have an interpretation, yielding nonsensical results for the implied persistence. Overall, the results suggest that excess sensitivity cannot account for the observed impulse responses.

Conversely, does the apparent resolution of the GR-puzzle showing up in table 4.4 have anything to do with the permanence of shocks at all? Note that the theory restricts the current account to be more sensitive to country-specific shocks only to the degree that they do have permanent effects. If the current account 'overshoots' investment even if shocks do not have a permanent effect, then table 4 would be meaningless. Also this

issue can be addressed, now be letting ρ go to zero. Then, from (4.27) above we get

$$\frac{s_{11}}{s_{12}} = \frac{\alpha_1}{\alpha_2}$$

This ratio of the adjustment coefficients gives us the 'shadow' impulse response of the current account and investment if country-specific shocks are completely transitory. Our estimates of α_1/α_2 are given in table (4.6): the results are encouraging - with the exception of the United States, the implied response is now still negative but smaller than unity in absolute value, in the case of Canada even positive. This verifies that it is indeed the fact that country-specific shocks have permanent components that leads the current account to react more sensitively than investment.

Putting things together, we find that near random-walk behaviour of country-specific shocks cannot account for the Glick and Rogoff puzzle when a model is used that restricts the data less strongly than the Glick and Rogoff model. Rather, by focussing on a reduced-form cointegrated VAR, we could show that the current account is actually more sensitive to country-specific shocks than is investment and that this result is in fact due to permanent components in country-specific shocks - as is predicted by the theory. We conclude, in the spirit of Quah (1990), that the GR-puzzle is likely to come about because estimated responses are an amalgam of responses to transitory and permanent shocks. We draw a conclusion similar to Quah's: univariate time series properties (i.e. the fact that TFP seems well described by a random-walk in a univariate context) should not be used as a basis for economic theorizing if the economic theory of interest involves several variables. We have proposed to focus on a few reduced-form implications of the theory and then to assess time-series properties in a dynamic system-framework. In the next subsection we will deal with the dynamic implications of the theory: impulse responses and the forecast performance of our models.

112

Hoffmann, Mathias (1999), International Macroeconomic Fluctuations, Capital Mobility and the Current Account: A cointegrated approach
European University Institute
DOI: 10.2870/45469

4.4.3   Dynamic Responses

The dynamic responses of the model are in line with what one would expect from the theory: Figures 3 to 8 provide plots of the dynamic response of the model for the G3 countries. The current account and investment react in different directions with respect to a country-specific shock. Both investment and the current account reach their permanent value after roughly five years. In the case of the U.S. and Japan, this means that the current account reverts to zero, which is an outcome of the cointegrating relationship in the model. As, in the estimation of the model for Germany, we have imposed a non-trivial cointegrating relationship between investment and the current account, there is no need in this model for the current account to revert to zero. Indeed, in the German case, country-specific shocks do have a pronounced permanent effect on the current account.

The response of the current account to global shocks is much less pronounced than to country-specific shocks. In the US case, the point estimate of the response is on average smaller than the response to the country-specific disturbance by a factor of ten. Similar results, even though with somewhat smaller factors, ensue for the other countries. It seems, that the imposition that the current account's period zero response to a global shock is zero is compatible with the data. In all three countries, however, the global shock has a noticeable impact on the permanent value of investment.

For Japan and the U.S. the responses to permanent and transitory shocks are largely unspectacular. The permanent shock has a sizeable impact on investment whereas the long-run response is zero for the current account. Only in the German case, the long-run response of the current account is roughly half of the investment response. To the degree that we believe that the cointegrating relationship between current account and investment reflects economic structure, this result tells us that permanent shocks in Germany (which over the sample period proved to be largely idiosyncratic), have huge leakage effects: the shock triggers increased investment but it also increases capital exports and hence leads to accelerated accumulation of foreign assets. Another notion is the one of export-led growth that is often referred to in the discussion about Germany's

113

postwar economic development (see e.g. Marin (1992)). We checked whether we could accept that the current account is weakly exogenous with respect to the parameters of investment. Indeed, this hypothesis could not be rejected. In effect this means that for Germany, innovations to the current account seem to represent permanent, country-specific shocks.

4.4.4 Forecast performance

Figures 7-12 display the results of a forecasting exercise. It is based on the following present value formula:

$$CA_t^p = -\sum_{l=1}^{\infty} R^{-l} \Delta I_{t+i|t}^c \qquad (4.32)$$

where $\Delta I_{t+i|t}^c$ represents the time t forecast of those changes in investment in time $t + i$ that are explained by country-specific shocks. Usually, in intertemporal optimization models, the current account is represented as the discounted sum of changes in net output, $NO_t = Y_t - C_t - G_t$ where G_t is government consumption. We deviate from this representation in this case and use investment as a proxy of net output. This allows us to stay in the framework of the econometric model we have used from the outset. As we will see, it seems a valid approach. The forecast performance of our model is very good and there is also a good rationale of why investment should be a good proxy of net output: models of balanced growth suggest that the great ratios, i.e. investment over output and consumption over output are stationary. Hence, changes in investment should be highly correlated with changes in output and we should not be too surprised to see the former predict the latter well.

In our VAR model, the predicted country-specific component of investment is given

by

$$\Delta I_{t+i|t}^c = \begin{bmatrix} 1, & 0 \end{bmatrix} \sum\nolimits_{l=i}^{\infty} C_l S \begin{bmatrix} e_{t-(l-i)}^c \\ 0 \end{bmatrix} \tag{4.33}$$

The values of $\Delta I_{t+i|t}^c$ are gained from this formula and then plugged into the above present-value relation in order to get CA_t^p. In figures 9-15, CA_t^p is then plotted together with the actual current account.

Overall, our models do a good job in tracking the current account dynamics. But also the order of magnitude of the swings in the current account is captured well in most cases. Even notoriously 'difficult' cases like Germany and the United States can be explained well by our models. The fit for France and Italy and also for Canada is very good. For Japan - based on a visual inspection of the plots - we get the ups and downs right but the variance is not quite precisely estimated. The UK remains the difficult case it usually is in the current account literature, the current account that is predicted by country-specific shocks alone is essentially flat. However,we calculated a correlation between the forecast and the observed current account of roughly 0.82, quite high vis-a-vis other studies (Gosh (1995) finds a correlation of 0.7 for the period 1960-88). Note that this result has been obtained by conditioning on the price of oil which does not figure in the models in the literature. As the country is a big oil exporter, its current account is likely to reflect the swings in the price of oil. To the degree that we consider oil price changes as global shocks, one would expect the British current account indeed to be better explained by global shocks rather than country-specific ones. Figure 16 shows the forecast of the current account, this time based on global rather than country-specific changes in investment. The forecast is certainly not good, but it is probably closer to the observed current account in terms of volatility than the forecast based on country-specific shocks.

Overall, the forecast performance of our models compares very well with that of earlier 'naive' approaches that do not take into account the distinction between country-specific

115

and global shocks. In some difficult cases like Germany and the US, our forecast is even much better. Even though it should be noted however, that we also obtained these improvements through conditioning on a set of exogenous variables, the models seem to fulfill the restriction imposed by economic theory, namely that only country-specific shocks drive the current account.

4.4.5 How country-specific are country-specific shocks?

Our discussion in the previous subsections documents a very good match between the theory and the data. However, we should recall that our identification procedure for country-specific shocks relied on the theory itself. We assumed that global shocks do not affect the current account in period zero. Certainly, this theoretical presumption is also backed by the results of Glick and Rogoff. Nonetheless, it would be nice to have an evaluation to know if we have really identified the right shocks. There is clearly no way in which we can evaluate a just-identifying assumption within each individual model. However, we have valuable information in the cross-section of countries we are investigating. The G7 countries account for two thirds of world output and they represent a fairly closed bloc in the world economy. It therefore seems reasonable to take these countries as a proxy of the 'rest of the world'. Country-specific shocks should then be uncorrelated across countries whereas we should find some correlation between the global shocks identified at the country level.

Table 4.7 gives the average correlation of each country's specific and global shocks with all other 6 countries. It also provides the standard errors of these correlations. The result is very encouraging: not only are global shocks much more highly correlated across countries than country-specific shocks, their correlation is also highly significant. On the other hand, country-specific shocks are on average not significantly correlated. The only exception is Canada, where both country-specific shocks and global shocks are on average significantly correlated with shocks in the rest of the world. Still, these results should provide some confidence that by and large we have indeed identified the right shocks.

116

4.5 Conclusion

The intertemporal approach to the current account is becoming increasingly standard in international macroeconomics. This theory makes very strong predictions about shocks that can be classified according to two criteria: persistence and country-specificity. The current account is supposed to respond only to the persistent but transitory component of shocks and this only to the degree that they are country-specific.

Little work has been done so far on classifying shocks along these lines and testing the predictions of the theory. The seminal paper by Glick and Rogoff (1995) is an exception. Whereas the structural estimation approach adopted by Glick and Rogoff allows us to understand in detail in which way the implied responses of investment and the current account depend on the persistence of country-specific shocks, the estimation itself relies on univariate evidence about the time-series properties of shocks, leading to estimates in which the relative sensitivity of the current-account and investment are at odds with the theory.

In this paper, we reverted to the more black-box approach of a structural VAR. Whereas this forces us to sacrifice some model structure, it puts us in a position to classify shocks to the current account and investment according to their persistence by exploiting cointegration information in the data. We identified country-specific shocks using the suggestions of the theory and the empirical results of Glick and Rogoff: global shocks do not have an effect on the current account. It then becomes possible to measure the persistence of country-specific shocks. We also derived a reduced-form analogue to the Glick-Rogoff result that the relative response of current account and investment is highly sensitive with respect to the persistence of country-specific shocks. In our estimates the puzzle encountered by Glick and Rogoff, i.e. that the relative response of investment vis-a-vis the current account is 2-4 times too strong, vanishes. As our results show the GR-puzzle is likely to have arisen because country-specific shocks have both important permanent and transitory components and therefore the impulse responses by Glick and Rogoff are likely to reflect an amalgam of responses to permanent and

117

transitory shocks. Our conclusion is that it is not possible to disentangle these permanent and transitory components unless the data are allowed to speak loudly and only some key restrictions are imposed from economic theory on the reduced form. In a more theoretical context, Quah (1990) has proposed the mechanism put forward in this paper as an explanation of the apparently excessively smooth behaviour of consumption vis-a-vis other macroeconomic aggregates, in particular output. Only if all shocks are permanent should consumption move one to one with permanent income. However, if economic agents distinguish between permanent and transitory shocks, consumption will on average be much smoother than output.

In this paper, we empirically explore the open-economy analogue of the excess-smoothness puzzle: if country-specific shocks have permanent and transitory components then the current account can be extremely sensitive to permanent shocks while at the same time being imperfectly correlated with investment.

Finally, we have exploited our approach to forecast the current account based only on the country-specific shocks. The forecast performance compares very well with models that are less restricted than ours. This provides evidence that the current account is indeed driven mainly by country-specific shocks. Even in the case of the United Kingdom we can gain some ground. Using investment as a proxy of net output and conditioning on oil prices, we can not only achieve a high correlation between the actual and the forecasted current account but also emulate the actual current account variance.

118

Bibliography

[1] Beveridge, S. and Nelson, C. R., (1981), 'A New Approach to the Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to the Measurement of the Business Cycle,' *Journal of Monetary Economics* 7: 151-74.

[2] Boltho, A. (1996), 'Was Japanese Growth Export-Led?', *Oxford Economic Papers*: 415-32.

[3] Campbell, J. Y, and Shiller, R.J., (1987), ' Cointegration and tests of present value models,' *Journal of Political Economy*, 95: 1062-1088.

[4] Feldstein, M. and Horioka, C., (1980), 'Domestic Saving and International Capital Flows', *Economic Journal* 90: 314-329.

[5] Ghosh, A., (1995), 'Capital mobility amongst the major industrialized countries: Too little or too much?,' *Economic Journal* 105: 107-128

[6] Glick, R. and Rogoff, K., (1995), 'Global versus country-specific productivity shocks and the current account,' *Journal of Monetary Economics* 35: 159-92.

[7] Gordon, Robert J., (1993), *Macroeconomics*, Addison-Wesley. Data set to be found at *http://www.nber.org/pub/gordon.*

119

[8] Harbo, I., Johansen, S., Nielsen, B. and Rahbek, A. (1998), 'Asymptotic Inference on Cointegrating Rank in Partial Systems', *Journal of Business and Economics Statistics* 16: 388-99.

[9] Hoffmann, M. (1998), 'Long Run Capital Flows and the Feldstein-Horioka Puzzle: A New Measure of International Capital Mobility and Some Historical Evidence from Great Britain and the United States', European University Institute, Department of Economics Working Paper no. 98/30.

[10] Johansen, S., (1988), 'Statistical Analysis of Cointegrating Vectors', *Journal of Economic Dynamics and Control*, 12:231-54.

[11] Marin, D. (1992), 'Is the Export-Led Growth Hypothesis valid for industrialized countries?', *Review of Economics and Statistics* 74: 678-88.

[12] Obstfeld, M., (1986), 'Capital Mobility in the World Economy: theory and measurement', *Carnegie Rochester Conference Series on Public Policy* 24: 55-104.

[13] Obstfeld, M., (1995), 'International Capital Mobility in the 1990s' in Kenen, P. (ed.) *Understanding Interdependence*, Princeton University Press, 1995.

[14] Obstfeld, M. and Rogoff, K. (1995), 'The intertemporal approach to the current account,' NBER working paper 4893.

[15] Obstfeld, M. and Rogoff, K., (1996), *Foundations of International Macroeconomics*, MIT Press, Cambridge.

[16] Penati, A. and Dooley, M., (1984), 'Current Account Imbalances and Capital Formation in Industrial Countries 1949-81', *IMF Staff Papers* 31: 1-24.

[17] Quah, D. (1990), 'Permanent and transitory movements in labor income: an explanation for "Excess Smoothness" in Consumption', *Journal of Political Economy* 98: 449-75.

[18] Sachs, J, (1981), 'The Current Account and Macroeconomics adjustment in the 1970s,' *Brookings Papers on Economic Activity* 1:201-268.

[19] Sheffrin, S. and Woo, W. T. (1990), 'Present Value Tests of an Intertemporal Model of the Current Account' *Journal of International Economics* 29: 237-53.

[20] Taylor, A. M., (1996), 'International Capital Mobility in History: The Saving-Investment Relationship', NBER Discussion Paper Series, no. 5743

4.6   Mathematical Appendix

We can rearrange (4.27) to yield

$$\rho^2\left(\alpha_2^2\omega_{11} + \alpha_1^2\omega_{22}\right) - \alpha_2^2\omega_{11} = \left(\rho^2 - 1\right)2\alpha_1\alpha_2\phi\sqrt{\omega_{11}\omega_{22}} + \alpha_1^2\phi\omega_{22} \tag{4.34}$$

For simplicity, we redefine

$$A = \left(\alpha_2^2\omega_{11} + \alpha_1^2\omega_{22}\right) \tag{4.35}$$

$$B = 2\alpha_1\alpha_2\sqrt{\omega_{11}\omega_{22}} \tag{4.36}$$

$$C_1 = \alpha_2^2\omega_{11} \tag{4.37}$$

$$C_2 = \alpha_1^2\omega_{22} \tag{4.38}$$

121

Substituting and rearranging, we get

$$G(\rho) = \rho^2 \text{ and } F(\phi) = \frac{C_1 - \phi B + C_2 \phi^2}{A - \phi B}$$

and

$$G(\rho) - F(\phi) = 0 \tag{4.39}$$

By the implicit function theorem

$$\frac{\partial \phi}{\partial \rho} = \frac{2\rho(A - \phi B)^2}{(2C_2\phi - B)(A - \phi B) + B(C_1 - \phi B + C_2\phi^2)} \tag{4.40}$$

Letting $\rho = 1$ implies $\phi = \pm 1$ and hence, exploiting $A - B = (C_1 - B + C_2)$, we get the result

$$\frac{\partial \phi}{\partial \rho}_{|\rho=1} = \pm \frac{(A \mp B)}{C_2} = \left[1 \mp \frac{\alpha_2}{\alpha_1} \sqrt{\frac{\omega_{11}}{\omega_{22}}} \right]^2$$

The economically relevant case is $\phi = -1$, (investment and the current account are negatively correlated). So we get

$$\frac{\partial \phi}{\partial \rho}_{|\rho=1} = \frac{-(A + B)}{C_2} = \left[1 + \frac{\alpha_2}{\alpha_1} \sqrt{\frac{\omega_{11}}{\omega_{22}}} \right]^2$$

122

4.7 Tables and Figures

Table 4.1: Tests for cointegration

a) Johansen Trace statistic

	US	Japan	Germany	France	Italy	UK	Canada	90%	95%
$h = 0$	25.4	25.06	20.78	19.58	19.57	14.29	13.82	15.58	17.84
$h = 1$	4.85	0.02	0.12	2.84	1.137	2.13	3.07	6.69	8.08

b) Johansen Maximum Eigenvalue statistic

	US	Japan	Germany	France	Italy	UK	Canada	90%	95%
$h = 0$	20.55	25.04	20.66	16.73	18.43	12.16	10.74	12.78	14.6
$h = 1$	4.85	0.02	0.12	2.84	1.137	2.13	3.07	6.69	8.08

The tests were performed on VAR(2)-models with an unrestricted constant.
The models for the US, Japan, Germany, the UK and Canada.
included one weakly exogenous regressor. Critical values for the trace test,
following table 3 in Harbo et. al. in this case are 10.4 (12.3) at 90 (95)%.

Table 4.2: Estimates of the cointegrating vector

Estimate of $\beta = \begin{bmatrix} 1 & \beta_2 \end{bmatrix}$ and test of $H_0 : \beta_2 = 0$

	US	Japan	Germany	France	Italy	UK	Canada
β_2	-0.2535	0.0174	**-0.619**	-0.002278	-0.005234	0.1728	0.0883
LR-test	1.91	0.2482	12.8	0.005503	0.0113	1.04	2.27
P-value	0.17	0.62	0.0003	0.94	0.92	0.6922	0.13

Table 4.3: persistence of country-specific shocks

	US	Japan	Germany	France	Italy	UK	Canada	Average
ρ^2	0.1702	0.3827	0.8656	0.1025	0.0474	0.0454	0.0329	0.2352

124

Table 4.4: Estimates of the Choleski factors

Coefficients	US	Japan	Germany	France	Italy	UK	Canada	Average
s_{11}	15.4	1.659	12.18	27.06	7.716	3.6386	3.498	-
s_{21}	-30.15	-1.012	-6.649	-21.03	-8.906	-3.5470	-2.361	-
s_{22}	42.25	2.858	14.19	38.95	6.318	3.8098	4.528	-
s_{11}/s_{21}	-0.511	-1.639	-1.831	-1.287	-0.8664	-1.0258	-1.482	-1.2345

Table 4.5: GR-responses and their implied persistence

	US	Japan	Germany	France	Italy	UK	Canada	Avg.
G&R	-0.2727	-0.3023	-0.4255	-0.275	-0.5	-1.02	-0.4884	-0.4692
$\chi(1)$	-0.2968	-0.5472	-0.777	-0.6113	-0.7066	-0.699	-0.7653	-0.629
implied ρ	0.9448	0.8962	0.9674	0.8249	0.9227	1.136	2392	0.9539[*)]
$\partial\chi(1)/\partial\rho$	0.436	2.358	10.77	1.92	2.672	2.369	0.0001	3.4210[*)]

[*)] not including Canada

Table 4.6: Implied response at $\rho = 0$.

	US	Japan	Germany	France	Italy	UK	Canada	Average
α_1/α_2	-1.4	-0.5086	-0.2854	-0.7915	-0.748	-0.8312	0.756	-0.544

Table 4.7: Cross-country correlations of structural shocks

a) country-specific shocks (e^c)

	US	Japan	Germany	France	Italy	UK	Canada
avg. correlation	0.00731	0.1112	-0.04689	0.08786	0.1378	0.02057	**-0.1623**
standard dev.	0.02016	0.07102	0.04186	0.09176	0.07629	0.01986	0.03637

b) global shocks (e^w)

	US	Japan	Germany	France	Italy	UK	Canada
avg. correlation	**0.2136**	**0.2842**	**0.217**	**0.3556**	**0.2025**	**0.2802**	**0.1658**
standard dev.	0.03877	0.04214	0.03662	0.01577	0.0401	0.009901	0.06059
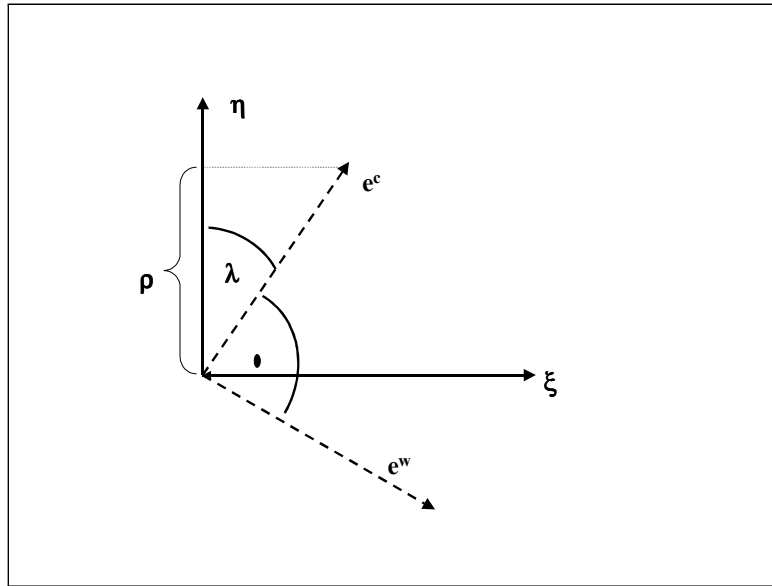
Figure 4-1: The geometry of global and country-specific shocks
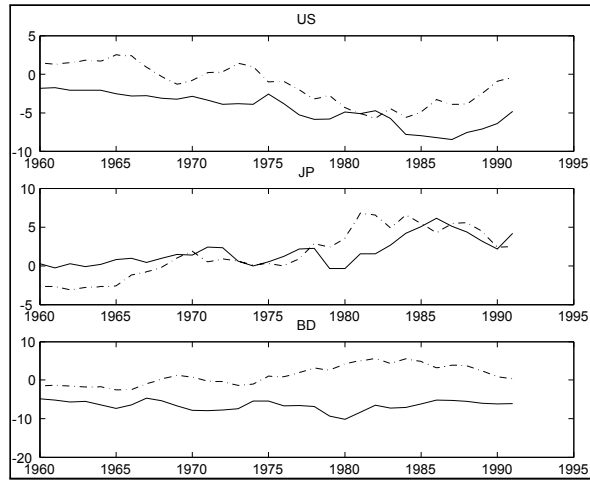
Figure 4-2: G 3 - interest rate differential (dashed) vs. cointegrating residuals.
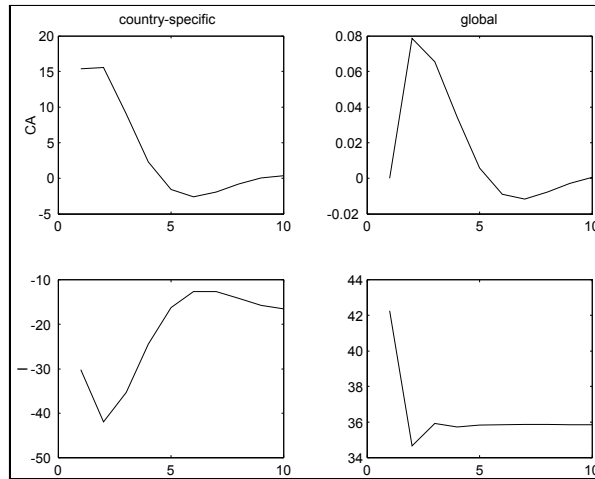
Figure 4-3: US - impulse responses by country_specificity

127
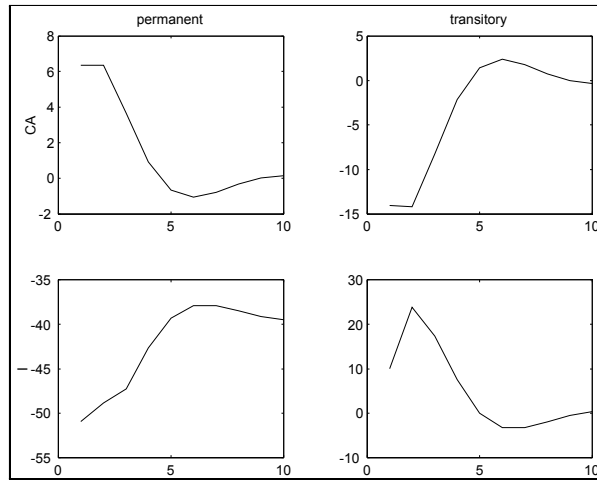
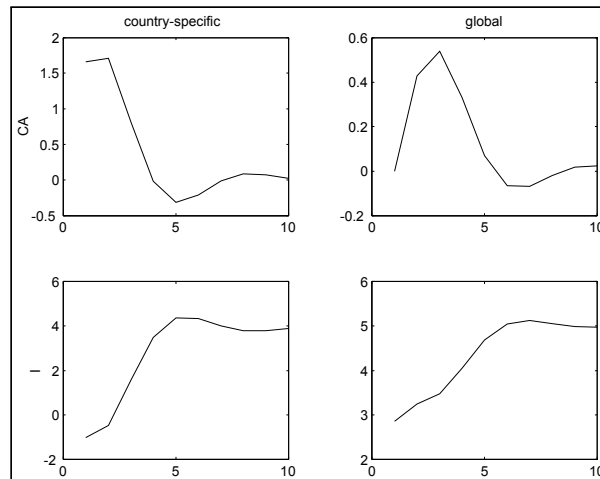Figure 4-4: US - impulse responses by persistence

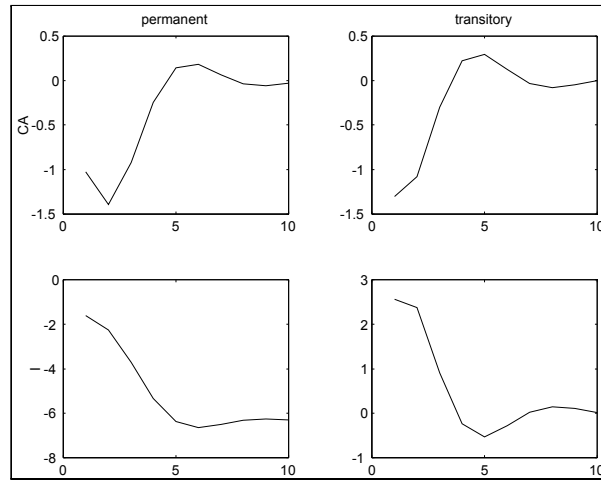Figure 4-5: Japan - impulse responses by country -specificity

128

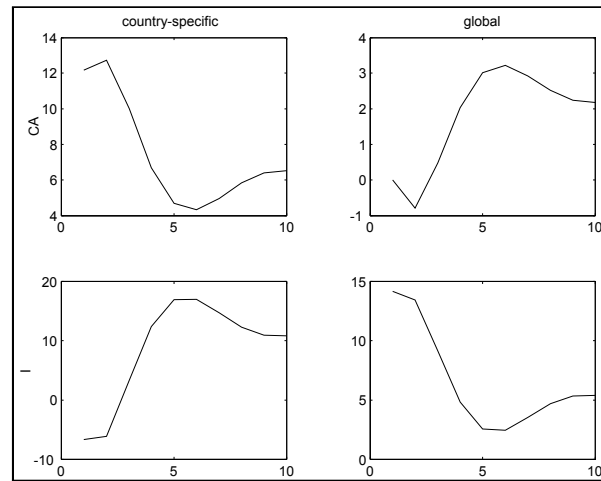Figure 4-6: Japan - impulse responses by persistence

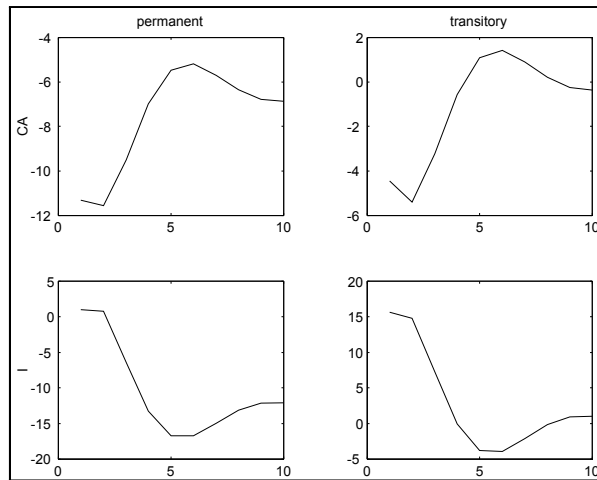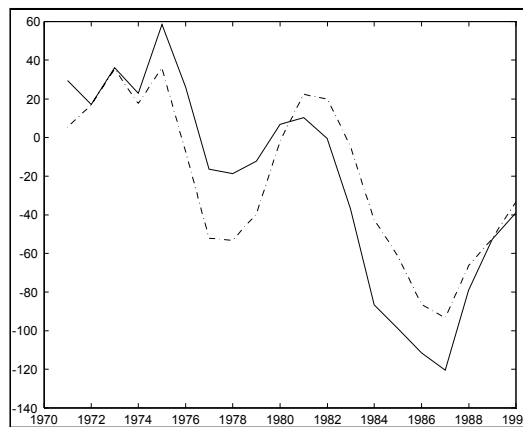

Figure 4-7: Germany - impulse responses by country-specificity

129

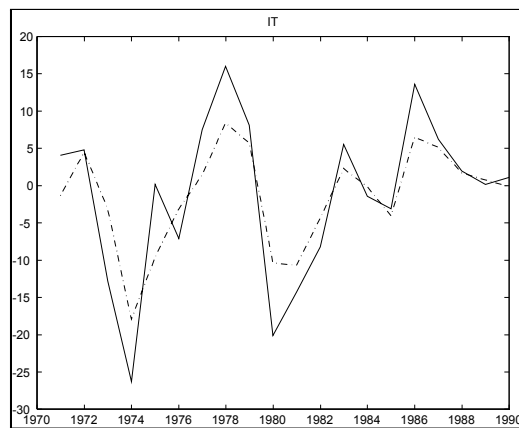Figure 4-8: Germany - impulse response by persistence

Figure 4-9: Actual and forecasted (dashed line) US current account

130
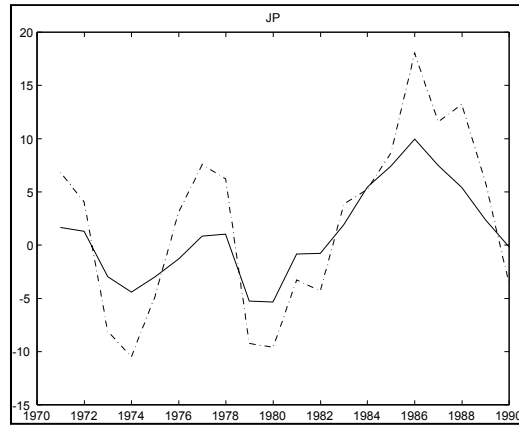
Figure 4-10: Actual and forecasted (dashed line) Japanese current account
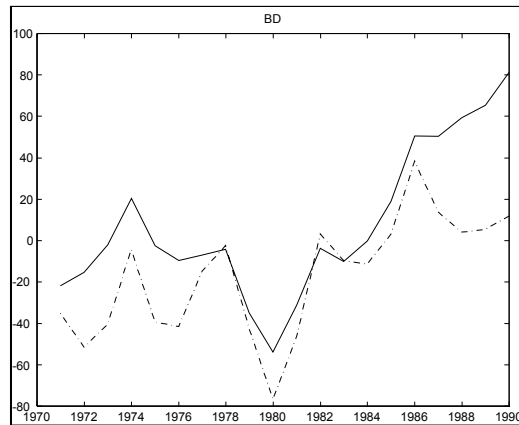
Figure 4-11: Actual and forecasted (dashed line) German current account

131
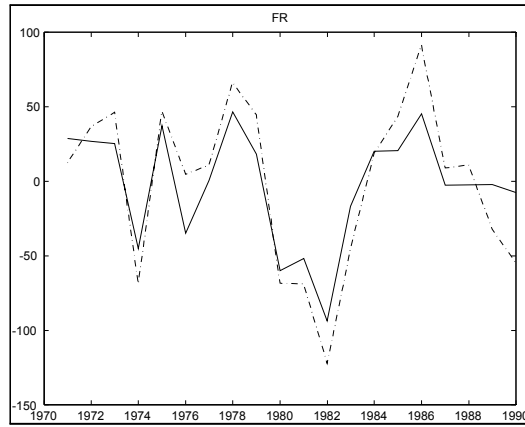
Figure 4-12: Actual and forecasted (dashed line) French current account

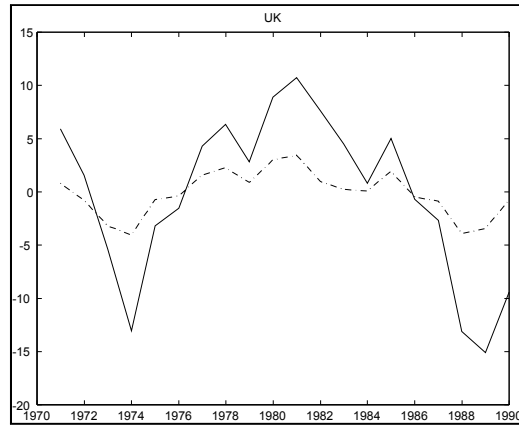Figure 4-13: Actual and forecasted (dashed line) Italian current account

132

Figure 4-14: Actual and forecasted (dashed line) UK current account
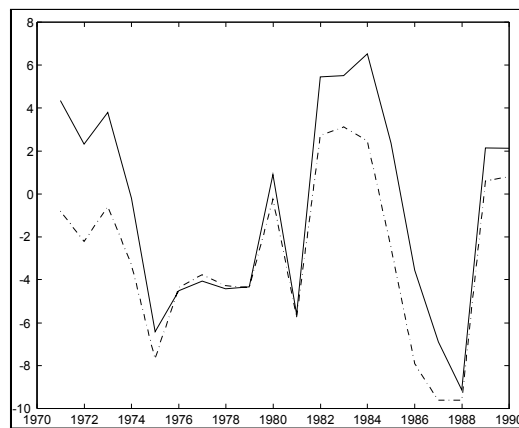

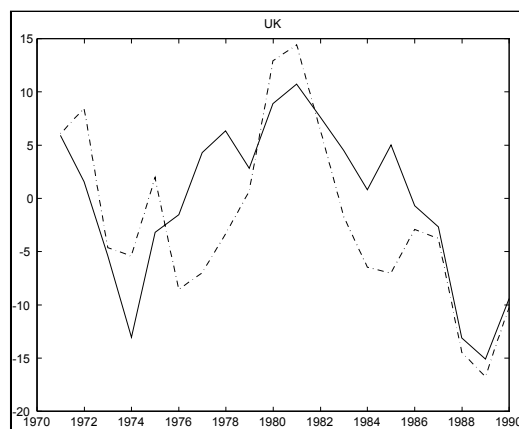Figure 4-15: Actual and forecasted (dashed line) Canadian current account

133

Figure 4-16: forecast of the UK current account based on both country-specific and global shocks

134