ESSAYS ON THE TRANSMISSION MECHANISM OF MONETARY POLICY

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Abstract of Chapter 1

In this chapter, the monetary version of the sticky price intertemporal model of Obstfeld and Rogoff (1995, 1996), in which unexpected and expansive monetary shocks unambiguously generate a permanent nominal exchange rate depreciation and a temporary current account surplus, is outlined. After discussing some extensions of the basic model and verifying the lack of robustness of the main theoretical predictions to the introduction of alternative assumptions, the chapter provides an empirical investigation of the role of nominal disturbances for current account and real exchange rate fluctuations within a structural VAR approach for 15 OECD countries over the period 1979-1998, using the long-run restriction identification scheme suggested by Clarida and Galì (1994). The main empirical findings suggest that that nominal shocks tend to have a significant role in generating temporary current account surpluses and that these effects are proportional to the degree of openness of the country.

Abstract of Chapter 2

Housing systems, as a major sector of industrialised economies, might have profound effects on the transmission mechanism of a monetary shock. Despite a progressive convergence, however, EU countries still differ significantly in their housing and credit market institutions. This chapter provides a theoretical discussion of the 'housing market' channels of the monetary transmission mechanism (MTM) and offers some evidence on institutional differences across EU countries. Using recursive and semi-structural VARs, the role of house prices in the MTM is then assessed in eight European countries over the pre-EMU period. Results show a different degree of sensitivity of house prices, partly consistent with the institutional features of the European housing systems. The importance of these policy-induced changes in house prices to the transmission of monetary shocks to private consumption are then investigated. The chapter provides some support for the view that the house price channel may be an important source of MTM to consumption in those economies where housing and mortgage markets are relatively more developed and competitive.

Abstract of Chapter 3

This chapter extends the existing cross-country housing empirical literature focusing on the main fundamental factors affecting house price dynamics in a number of ways. First, through the implementation of seemingly unrelated regression (SUR) techniques and heterogeneous panel estimation methods, it is shown that European house prices are asymmetrically affected by real and financial variables. Subsequently, using a recent dataset, which collects quarterly information on housing and mortgage markets of EU countries, separate house price equations are estimated within an unrestricted error correction mechanism (ECM) approach for eleven European economies over the period 1980-2001. Results show that European house prices are driven by similar factors, but that their relative importance differs very significantly across countries. In particular, while real income is the single most important determinant of real house prices, financial effects play a relatively more important role in those countries that experienced a higher degree of financial liberalisation.

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INTRODUCTION

MOTIVATIONS AND THESIS STRUCTURE

1. Background

Over the last few decades, many economists have been focusing on the question of whether and how monetary policy affects the level of activity in the economy. Whilst a widespread consensus has been reached on the short to medium-run effects of monetary policy shocks on the real activity and long-run effects on the price level, the debate over the relative importance of the channels of monetary transmission mechanism (MTM) is still open.

Economists identify a number of different channels through which a monetary policy change can influence the real activity. In particular, they have developed two main views which are not mutually exclusive: the 'money view' and the 'credit view.' The former refers to the traditional understanding of monetary transmission to real activity based on the standard IS-LM model, in which money supply and interest rate movements directly affect the level of consumption and investment expenditure through the traditional cost of capital, substitution, income, and asset price channels. The credit view, on the other hand, emphasises the importance of financial market imperfections and recognises the special role played by financial intermediaries in transmitting monetary impulses to output and inflation.

In the light of the recent findings of the monetary theoretical literature and the dramatic developments in the economic and financial structure of the industrialised economies, new insights and important questions have recently attracted the attention of academics and policy-makers in many industrialised countries. This thesis aims to make a contribution to this discussion by developing three empirical studies, each of which focuses upon either specific channels of the monetary transmission mechanism, or aspects which have direct implications for the formulation of appropriate policy responses.

During the last few decades, international macroeconomic research has gone through significant developments. Amongst the most important of these is the introduction of new theoretical models, based upon dynamic general equilibrium frameworks combining microfoundations, intertemporal decisions and market imperfections in the form of nominal rigidities and monopolistic competition. This new line of research (also known as new open economy macroeconomics), which began with the publication of the *Redux* article by Obstfeld and Rogoff (1995), has attracted the interest of many international macroeconomists, who have extended the original setup in a number of directions to study crucial aspects such as optimal monetary policy rules, the benefits of international policy coordination and the effects of structural shocks on the economy. Key variables in this class of models are the exchange rate and the current account, which play a primary role in the international monetary transmission mechanism in open economies. The theoretical implications for the dynamics and the volatility of the latter variables in response to nominal disturbances, however, are mixed and the predictions of new open economy models appear to be sensitive to the specification of the microfoundations. This implies that it is difficult to carry out welfare analysis and discuss policy evaluation if there is no agreement on the correct or 'generally accepted' theoretical framework. Although the new literature has developed very rapidly, relatively little empirical work has been carried out to test the accuracy of the theoretical predictions. Chapter 1 of this dissertation makes a step in this direction by focusing upon the dynamics of the current account, one of the fundamental variables in the international monetary transmission mechanism in the new open economy literature.

The other major topic developed in this dissertation refers to a different aspect, which, nevertheless, has over the last few years increasingly attracted the attention of many academics and monetary policy authorities: the housing system. This considerable interest is justified by a number of reasons. First, housing represents more than half of the net wealth of the private sector in most industrialised economies. Second, home ownership is distributed much more equally than financial assets and represents one of the main transactions in the lives of most individuals. Third, housing has a crucial collateral role in the household lending sector, and fluctuations in house prices could have a significant impact in prompting financial cycles as well as implications for financial stability. As a result, several international institutions and central banks have become increasingly interested in the role of residential asset prices in the transmission mechanism of monetary policy and in the identification of the sources and the nature of the shocks driving house prices. This is further justified by the financial deregulation process of the 1980s and 1990s, which, in reshaping the European housing financing systems, made residential asset prices more and more important in affecting the aggregate demand. At the same time, however, the housing and mortgage markets have remained very segmented across countries, rendering them a potential source of real asymmetries. Chapters 2 and 3 of this dissertation provide a detailed study of the European housing and mortgage markets over the recent decades. In particular, the former focuses on the role of house price in the transmission of exogenous interest-rate changes in selected European countries during the pre-EMU period, whereas the latter studies the sources and nature of the shocks driving residential prices in the European Union (EU).

In order to better appreciate the value-added of this thesis, the next section provides a summary of the motivations and techniques employed in each chapter and the contributions they make to the relevant literature. The final section briefly outlines some comments upon the limitations of policy implications that could be drawn from this dissertation.

2. *Thesis Structure*

Chapter 1:
Nominal Shocks and the Current Account: a SVAR Analysis of 15 OECD Countries

This part focuses upon the role of nominal disturbances for exchange rate and current account fluctuations, which are fundamental variables in the international monetary transmission mechanism of the new open economy macroeconomics (NOEM). Differently from the previous intertemporal literature characterised by flexible-price and non-monetary economies, the latter theoretical models introduce market imperfections (namely, price stickiness and monopolistic competition) providing new insights in the role of monetary disturbances as tool of economic stabilisation in open economies, and challenging the traditional Keynesian aggregative models. Since the seminal work of Obstfeld and Rogoff, however, there has been a huge proliferation of models which, closely following the original setup, have extended and relaxed some of key assumptions of the *Redux* framework and showed that many theoretical predictions are not robust to alternative specifications of the microfoundations. Amongst the most controversial, the NOEM seems to provide contrasting theoretical predictions on the effects of monetary disturbances on the current account and its role in the international monetary transmission mechanism.

This chapter directly tackles this deficiency and attempts to answer some questions which have fundamental policy-implications. In particular, are nominal disturbances a significant factor driving current account fluctuations in open economies? Which theoretical framework is empirically more relevant? That is, is it possible to distinguish between those classes of models predicting a current account surplus in response to an unexpected monetary shock, and those in which a current account neutrality or deficit might occur? Is the external account an essential adjustment variable in the international monetary transmission mechanism? If so, does this channel have an increasing potency according to the degree of openness of a country?

In order to reply to these questions, the previous empirical literature (mainly limited to a small subset of open economies) is extended, using Structural Vector Autoregression (SVAR) models which are applied separately to fifteen OECD countries over the pre-EMU period. The identification of the structural forces driving the variables of the systems is achieved using long-run neutrality restrictions as originally pioneered by Blanchard and Quah (1989) and later extended in Clarida and Galì (1994).

The empirical results will consist of impulse-response simulations and variance decompositions, which are used to summarise the dynamic relationships between the variables and structural shocks in the estimated vector autoregressive models.

Chapter 2:
Monetary Policy Shocks and the Role of House Prices across European Countries

The years preceding the introduction of the Euro have been characterised by a growing concern regarding the potential problems associated with the conduct of a common monetary policy in an area characterised by heterogeneous members. Not only the possibility of idiosyncratic shocks, but also large differences in economic and financial structures among the EMU economies could represent a potential matter of concern for the European Central Bank (ECB). The latter aspects are particularly relevant because they are critical factors in affecting the magnitude and the timing of the monetary transmission mechanism (MTM).

Some authors argue that the regime switch created with the institution of the ECB renders any attempt to assess real asymmetries and to study the MTM using pre-EMU data inappropriate. However, it is doubtful that this institutional change brought about behavioural changes in a sharp and discontinuous fashion. As a result, many academics have been trying to compare and analyse the effects of a monetary policy shock across countries based on the previous regimes, with the aim of throwing some light on the country-specific MTM. However, few of them attempt to identify the importance of specific transmission channels. This chapter focuses upon a relatively unexplored MTM working through house price fluctuations.

Recently, many academics and policy-makers have stressed the role that real assets prices could play for the real economy and the MTM. Besides the reasons outline in the introduction, the financial deregulation process of the 1980s and the resulting rising competitive pressure of the credit market have made house prices increasingly significant for the economic activity. Although most economists would agree with the latter statements, there is still little empirical work assessing the quantitative importance of such a 'house-price-related' channel.

In order to answer this question, chapter 2 investigates the role of residential asset prices in the MTM in some European countries during the pre-EMU period and provides some qualitative and quantitative evidence on the link between unanticipated short-term interest rate changes and residential prices, and between house prices and household consumption (non-housing) expenditure. To allow for a better interpretation of the empirical results, the chapter also provides a descriptive analysis of the main features of the housing system across countries and identifies those characteristics which might affect the role of these assets in the MTM. Similar to the vast majority of the previous empirical literature, the econometric approach employed is the use of Structural Vector Autoregression models, in which unanticipated changes in the monetary authorities' actions are identified through short-run restrictions between the variables of the systems, as originally suggested in Sims (1980, 1986) and Bernanke (1986). For each country under investigation, the effects of house prices to short-term interest rate changes are estimated through the use of impulse-response functions and variance decomposition analyses. The amplifying importance of these policy-induced-house price effects for household consumption decisions is then assessed through the implementation of alternative counterfactual simulation exercises.

Chapter 3:

House price Dynamic in Europe: a Cross-Country Empirical Analysis

This chapter closely follows the previous contribution and focuses upon the study of the macroeconomic factors driving residential asset prices in most of the European countries using a newly constructed dataset collecting housing and mortgage market information over the recent decades. In particular, the results of the existing international housing literature are extended in number of ways, taking into account the high degree of heterogeneity, which seems to characterise the national housing markets across the European Union (EU) economies and, at the same time, using a common methodology to enhance cross-country comparability.

The first part of the chapter provides an extensive overview of the house price dynamics, with particular emphasis upon the degree of volatility, and the size and timing of the booms and busts over the last two decades. The main developments which occurred in the mortgage markets during the financial deregulation process of the 1980s and 1990s and specific individual features related to the housing tax treatment, transaction costs, and mortgage contracts are then discussed, making use of recent information collected from a number of national and international sources. The above sections display a representation of the national housing and mortgage markets which motivate the following empirical results.

In the second part of the chapter, the previous international empirical literature, mainly based upon pooled estimation techniques, is extended by implementing alternative dynamic heterogeneous panels techniques, namely Seemingly Unrelated Regression (SUR) methods and Pooled Mean Group (PMG) estimators, to account for slope-heterogeneity of the parameters associated with the short and long-run factors affecting real house prices in the EU countries. After formally testing for and rejecting the homogeneity of the parameters, the subsequent sections provide the results of alternative (but similar in their specification) national house price equations. In order to capture the short and long-run factors affecting house prices, traditional unrestricted Error Correction Mechanism (ECM) models are employed separately for each of the economies under investigation.

3. *Preliminary Remarks*

Before proceeding, it is worth pointing out some of the issues which limit the policy implications which can be drawn from this dissertation. First of all, most of the empirical estimates are based on the analysis of data before the implementation of the euro. For instance, the results of the first and second chapter focus on the pre-1999 period in which intra-EMU exchange rate changes were playing a role in the MTM and in the reaction functions of the national monetary authorities. Secondly, over the last few years capital markets, the banking (and, in particular, the mortgage) sector, expectations and credibility of the central bank authorities went through significant changes. Accordingly, it is not possible to clearly assess whether or not these developments have affected the monetary policy transmission mechanisms and the relative importance of specific channels in any dramatic manner. As a result, it is hard to draw any uncontroversial conclusions on how the MTM currently works across countries.

Additional problems faced in the study of the housing markets across countries are related to quality of the data used in the second and third chapter. Although great effort has been made to collect comparable and reliable indices across countries, house price data come from a variety of public and/or private sources with subsequent differences in their geographical coverage, the coverage of the types of dwellings, quality adjustment, etc. As a result, although they represent the best information currently available, the role of these methodological differences in the estimation results could be important. Nevertheless, the lack of more reliable and comparable statistics and a relatively limited presence of cross-country empirical studies make the results of this thesis particularly informative.

CHAPTER 1

NOMINAL SHOCKS AND THE CURRENT ACCOUNT: A STRUCTURAL VAR ANALYSIS OF 15 OECD COUNTRIES

1. Introduction

Over the last two decades, many researchers have tried to provide an analytical framework which could be a superior alternative to the Mundell-Fleming-Dornbusch (MFD) model. Although having a long history in open economy macroeconomics, the latter approach has important limitations. First, the MFD model does not account for intertemporal decisions of individuals. Second, the saving-investment and money-demand relationships are introduced in an *ad hoc* manner and there is no explicit theory of price setting (Devereux, 1997). To this regard, a number of studies published in the early 1980s moved towards the so-called 'intertemporal approach,' which is based on micro-founded dynamic optimising models, where preferences, technology and capital market access are directly included. Central to this new framework is the role of international asset markets in allowing economies to trade consumption goods over time by borrowing from and lending to each other (Obstfeld and Rogoff, 1996). This new framework represents a fundamental tool not only in explaining current account imbalances over time, but also studying the role of the 'intertemporal trade' in the propagation of business cycles and the effects of economic shocks.[1]

In particular, as emphasised in Obstfeld and Rogoff (1995a), the intertemporal approach to the current account represents a synthesis of the 'absorption approach' by recognising that private saving and investment decisions as well as government decisions are the result of forward-looking expectations, and the 'elasticity approach' where relative international prices and the price elasticities of demand and supply are the central determinants of the net export balance.

Whereas many academics addressed the intertemporal analysis of the current account, focusing on flexible-price and non-monetary economies, in which real (e.g. fiscal and technology) disturbances are the main driving factors (see Glick and Rogoff, 1995 and Obstfeld and Rogoff, 1995a for a survey), only recently, following the partial failure of this class of models in explaining the high volatility of the exchange rate and the current account, several economists have introduced a role for monetary shocks. The *Redux* model by Obstfeld and Rogoff (1995, 1996) (henceforth OR) is without doubt considered to be the precursor of this literature. By introducing market imperfections (namely, price stickiness and monopolistic competition) in an intertemporal optimising open economy framework, their model shows that, under floating exchange rates, monetary shocks unambiguously affect consumption, income and exchange rates, and that current accounts have an explicit role in the international monetary transmission mechanism.

[1] The behaviour of the current account is also largely studied for its role in currency and financial crises in emerging countries. See Edwards (2001) for a recent empirical investigation.

Since the publication of the *Redux* model there has been a phenomenal interest in the so-called 'New Open Economy Macroeconomics' (NOEM) and the subsequent literature extended and developed the OR framework in several directions, providing new theoretical insights and a number of testable predictions (see Lane, 2001 and Sarno, 2000 for a comprehensive survey). For instance, several researchers moved away from the purchasing power parity (PPP) assumption of the Obstfeld-Rogoff formulation, introducing international market segmentation in the form of pricing to market (PTM) behaviour. The latter dramatically reduces the magnitude of (or even eliminates) the current account response to nominal shocks, with subsequent implications on the monetary transmission mechanism (Betts and Devereoux, 1996, 1999, 2000). Other authors developed generalisations of the model, introducing different degrees of elasticity of substitution between home and foreign goods which lead expansive monetary shocks to generate deficits of the current account (Lombardo, 2001, 2002 and Tille, 2001, amongst others). Finally, due to the specific non-stationary properties of the *Redux* model (see Cavallo and Ghironi, 2002 for a discussion), some academics completely de-emphasised the role of external imbalances and, to simplify the analysis, neglected current account responses to shocks on the ground that, at a quantitative level, they play a marginal role in the monetary transmission mechanism (Lane, 2002). While the latter assumption is a good approximation for relatively closed economies, foreign trade and asset exchange have a central role for transferring resources over time in open countries.

While theoretical work in the NOEM literature has developed very rapidly, surprisingly, few empirical studies have been carried out to test the accuracy of the main predictions of these models and the degree to which such models can be trusted for policy analysis. This chapter aims to throw some light on this direction by focusing on one of the key variables for open economy macroeconomics, namely the current account. In particular, extending the previous empirical studies mainly based on the G7, this chapter provides the results of two structural VAR models estimated separately for 15 OECD economies over the pre-EMU period. The structural shocks driving the system dynamics are identified through the impositions of long-run neutrality restrictions as originally advanced in Blanchard and Quah (1989). This analysis has two main objectives. Firstly, and differently from the previous empirical literature which mainly focuses on real disturbances (Glick and Rogoff, 1995), the chapter directly investigates the contribution of monetary disturbances to explain fluctuations of the current account. Secondly, it tries to empirically validate the implications of those classes of models predicting an external surplus versus those ones suggesting a neutrality or deficit in response to expansive nominal shocks.

The rest of the chapter is organised as follows. Section 2 illustrates the monetary version of the *Redux* model which motivates the empirical analysis. Section 3 briefly summarises some extensions of the original framework in the form of pricing to market and different degrees of elasticity of substitution between home and foreign goods, which dramatically affect the main predictions of the OR model. Section 4 provides an overview of some recent empirical attempts to test the role of monetary shocks in open economies with a particular emphasis on the effects of nominal shocks on exchange rates and external imbalances. Section 5 contains a description of the data and the econometric approach, and a discussion of the specifications and the corresponding estimation results. Finally, Section 6 concludes.

2. *A Two-Country General Equilibrium Model of International Monetary Policy Transmission (Obstfeld-Rogoff Redux Model, 1995, 1996)*

For several years, the traditional aggregative Keynesian models with sticky prices and the flexible-price intertemporal neoclassical models have represented the main way of modelling the economy. Both of them, however, present weaknesses. The former ones, although allowing for nominal rigidities, have many shortcomings linked to the lack of micro-foundations for intertemporal choice. In this regard, this class of models says very little about current accounts dynamics,[2] and, in general, does not provide a clear description of the intertemporal transmission mechanism of monetary shocks. Moreover, they do not allow for welfare and normative analysis, which enormously limit the possibility of formulating policy prescriptions. On the other hand, the neoclassical models, while embodying many of these central issues, are based on the hypothesis of flexible prices and perfectly competitive markets.

As pointed out in Walsh (1998), in order to create a bridge between these two approaches, Obstfeld and Rogoff (1995, 1996) develop a perfect foresight two-country general equilibrium model, which combines three fundamental blocks:

1. the first block is based on the intertemporal decisions by individual agents, where foreign trade and asset exchange allow for effects of transferring resources over time;
2. the second block relies on monopolistic competition in the goods market, which rigorously justifies the Keynesian assumption that output is demand-determined in the short run;

[2] The new intertemporal approach contrasts with the Keynesian models in that net exports are only determined by current relative income levels and net foreign interest payments are generally ignored. As a result, current account imbalances are not formally modelled.

3. the third block contemplates the presence of sticky prices, which, together with the monopolistic competition assumption, allows for real effects and non-neutrality of monetary policy actions.

In introducing nominal rigidities without abandoning the insights of modern intertemporal economics, they also address important issues in international finance like the volatility of the exchange rates, as well as the choice of exchange rate regimes.

What follows provides a detailed description of a simplified version of the original model, which mainly focuses on the international transmission mechanism of monetary policy on the nominal exchange rate, income, consumption and current account dynamics. The section below starts with an outline of the model with flexible prices. The subsequent part introduces preset nominal prices, and shows how monetary policy actions become a powerful stabilisation instrument.

2.1 The Two-Country Model with Flexible Prices

The main assumptions and features of the model can be summarized in the following components:[3]

2.1.1 Household's Preferences, Technology and Market Structure

The world is populated by a continuum of individual consumer-producers, indexed by $z \in [0, 1]$, each of whom produces a single differentiated perishable good, also indexed by z. The home country consists of producers on the interval $[0, n]$, and the remaining

[3] The content and the structure of this outline draws heavily from the original paper of Obstfeld and Rogoff (1995) and its monetary version presented in chapter 10 of Obstfeld and Rogoff (1996). In their framework, by excluding public expenditure, the main qualitative results with respect to the effects of a monetary shock are not affected.

agents $z \in (n, 1]$ reside in the foreign country. Thus, n provides an index of the relative size of the two countries. Foreign variables will be denoted by a superscript asterisk (*).

Individuals everywhere in the world have identical preferences defined over a consumption index, real money balances, and effort used in production. In particular, the intertemporal utility function of a typical home agent j depends positively on consumption (C) and real balances (M/P) and negatively on work effort, which is positively related to output (y). Its general formulation is given by:

$$U_t^j = \sum_{s=t}^{\infty} \beta^{s-t} \left[\frac{\sigma}{\sigma-1} C_s^{j\frac{\sigma-1}{\sigma}} + \frac{\chi}{1-\varepsilon}\left(\frac{M_s^j}{P_s}\right)^{1-\varepsilon} - \frac{k}{\mu} y_s(j)^{\mu} \right]$$

where $0 < \beta < 1$ is the subjective rate of discount, σ is the elasticity of intertemporal substitution, $\varepsilon > 0$ is a parameter inversely related to the elasticity of money demand, and μ is the elasticity of disutility from output. ? and k are positive 'scaling' parameters, giving the derived utility from real balance holding and leisure relative to the total utility of the representative consumer.

For analytical reasons, OR take into account the specific case:

$$(1) \qquad U_t = \sum_{s=t}^{\infty} \beta^{s-t} \left[\log C_s + \frac{\chi}{1-\varepsilon}\left(\frac{M_s}{P_s}\right)^{1-\varepsilon} - \frac{k}{2} y_s(z)^2 \right]$$

where it is assumed that all individuals have the same preferences, the elasticity of intertemporal substitution σ is equal to one and the elasticity of disutility from output μ is equal to two. Each component of the period utility function (1) is now discussed.

Assuming that all goods are traded and letting $c^j(z)$ be a home individual's consumption of product z, the variable C^j can be defined as a real consumption index (on which utility depends) and represented by the constant-elasticity-of-substitution (CES) function:

$$(2) \qquad C^j = \left[\int_0^1 c^j(z)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta}{\theta-1}}$$

where the elasticity of substitution between varieties $\theta > 1$[4]. The foreign consumption C^* is defined analogously.

The price deflator for nominal money balances is the consumption-based money price index corresponding to the real consumption index in eq. (2).[5] Letting $p(z)$ be the home-currency price of good z, the money price level in the home country can be defined as:

$$(3) \qquad P = \left[\int_0^1 p(z)^{1-\theta} dz \right]^{\frac{1}{1-\theta}}$$

Similarly, being $p^*(z)$ the foreign-currency price of good z, the foreign price index P^* can be written as:

$$(3.1) \qquad P^* = \left[\int_0^1 p^*(z)^{1-\theta} dz \right]^{\frac{1}{1-\theta}}$$

There are no impediments or costs to trade between the countries. Letting E be the nominal exchange rate, defined as the home-currency price of foreign currency, it is assumed that the law of one price (LOOP) holds for every good, so that:

$$(4) \qquad p(z) = Ep^*(z)$$

This implies that (3) and (3.1) can be rewritten also in the following form:

[4] This parameter will be the price elasticity of demand faced by each monopolist. This restriction has to be imposed because, since marginal revenue is negative when elasticity of demand is less than 1, $\theta > 1$ ensures an interior equilibrium with a positive level of output (Obstfeld and Rogoff, 1996, pp. 661).

[5] The price index is defined as the nominal expenditure of domestic money needed to purchase a unit of consumption C.

$$P=\left[\int_0^1 p(z)^{1-\theta}dz\right]^{\frac{1}{1-\theta}}=\left\{\int_0^n p(z)^{1-\theta}dz+\int_n^1\left[Ep*(z)\right]^{1-\theta}dz\right\}^{\frac{1}{1-\theta}}$$

and

$$P^*=\left[\int_0^1 p^*(z)^{1-\theta}dz\right]^{\frac{1}{1-\theta}}=\left\{\int_0^n [p(z)/E]^{1-\theta}dz+\int_n^1 p^*(z)^{1-\theta}dz\right\}^{\frac{1}{1-\theta}}$$

Since both countries' residents have the same preferences, equation (2) implies that home and foreign consumer price indexes are related by the purchasing power parity (PPP) condition:

(5) $\quad P=EP*$

There is no capital or investment in the model, but labour supply is endogenously determined. In fact period t output of good z, $y_t(z)$, is chosen in a manner that depends upon the marginal revenue of higher production, the disutility of effort, and the marginal utility of consumption. As a result, the level of output is endogenous. There is an integrated world capital market, in which both countries can borrow and lend. The only asset traded is a real bond, denominated in the composite consumption good C_t. Letting r_t be the real interest rate earned on bonds between t and $t+1$ and F_t and M_t the stocks of bonds and domestic money held by a home resident entering date $t+1$, individual z's period budget constraint (written in nominal terms) is given by:

(6) $\quad P_tF_t+M_t=P_t(1+r_{t-1})F_{t-1}+M_{t-1}+p_t(z)y_t(z)-P_tC_t-P_tT_t$

where $y(z)$ is the individual's output, for which agent z is the only producer, and T denotes lump-sum real taxes paid to the domestic government.

Since Ricardian equivalence holds in this model, it is assumed that the government runs a balanced budget during each period. In this version of the *Redux* framework, no government spending is considered and all seignorage revenues are distributed to the public in the form of transfers:

(7) $$0 = T_t + \frac{M_t - M_{t-1}}{P_t} \quad \text{or} \quad -P_t T_t = M_t - M_{t-1}$$

The same is true for the foreign country.

Given the CES consumption index in equation (2), it is possible to demonstrate that a home individual j's demand for good z in period t and the correspondent demand of a foreign individual are given by:[6]

$$c_t^j(z) = \left[\frac{p_t(z)}{P_t}\right]^{-\theta} C_t^j \qquad \text{and} \qquad c_t^{*j}(z) = \left[\frac{p_t^*(z)}{P_t^*}\right]^{-\theta} C_t^{*j}$$

Integrating demand for good z across all agents (that is, taking a population weighted average of home and foreign demands), and making use of eq. (2) and (4), which implies that $p(z)/P = p^*(z)/P^*$ for any good z, the constant-elasticity-of-substitution total world demand for good z is defined as:

(8) $$y_t^d(z) = \left[\frac{p_t(z)}{P_t}\right]^{-\theta} C_t^w$$

where world consumption C_t^w, which producers take as given, is:

(9) $$C_t^w = \int_0^n C^j dj + \int_n^1 C^{*j} dj = nC_t + (1-n)C_t^*$$

The latter relationship relies on the symmetry assumption (on the identical agents within each country), which is imposed to simplify the notation.

[6] See Appendix 1, Proof 1.

2.1.2 *Individual Maximisation and First-Order Conditions*

Using (8) to eliminate $p_t(z)$ from the period budget constraint (6) (implying that $p_t(z)y_t(z) = P_t y_t(z)^{(\theta-1)/\theta}(C_t^w)^{1/\theta}$) and dividing by P_t, the resulting expression can be substituted in the intertemporal utility function (1). Maximising the latter with respect to F_t, M_t and $y_t(z)$ leads to three fundamental first-order conditions (FOCs).[7] In particular, from the first FOC the corresponding standard consumption Euler equations are given by:

(10) $$C_{t+1} = \beta(1+r_t)C_t$$

and, for the foreign country:

(11) $$C_{t+1}^* = \beta(1+r_t)C_t^*$$

By using the definition of the home-currency nominal interest rate on date t, i_t, which, according to the Fischer parity equation, is defined as:

(12) $$(1+i_t) = \frac{P_{t+1}}{P_t}(1+r_t),$$

the second FOC leads to the money market equilibrium conditions for the home and foreign countries respectively:

(13) $$\frac{M_t}{P_t} = \left[\chi C_t\left(\frac{1+i_t}{i_t}\right)\right]^{1/\varepsilon}$$

and

(14) $$\frac{M_t^*}{P_t^*} = \left[\chi C_t^*\left(\frac{1+i_t^*}{i_t^*}\right)\right]^{1/\varepsilon}$$

Finally, the maximisation with respect to $y_t(z)$ leads to:

---

[7] See Appendix 1, Proof 2.

$$(15)\quad y_t(z)^{\frac{\theta-1}{\theta}} = \frac{\theta-1}{\theta k}(C_t^w)^{1/\theta}\frac{1}{C_t}$$

and

$$(16)\quad y_t^*(z)^{\frac{\theta-1}{\theta}} = \frac{\theta-1}{\theta k}(C_t^w)^{1/\theta}\frac{1}{C_t^*}\ .$$

2.1.3 Market-Clearing Conditions and the Initial Steady State

At the aggregate level, the domestic nominal money supply equals domestic nominal money demand in each country. Moreover, the global net foreign assets must be zero. This implies that:

$$(17)\quad nF_t + (1-n)F_t^* = 0$$

Under these bond-market-clearing conditions, dividing both sides of the home period budget constraint by the price level P_t and taking a population weighted-average of the resulting expression across home and foreign individuals, the following aggregate global goods-market clearing condition can be derived:

$$(18)\quad C_t^w = nC_t + (1-n)C_t^* = n\frac{p_t(h)}{P_t}y_t(h) + (1-n)\frac{p_t^*(f)}{P_t^*}y_t^*(f) = y_t^w$$

which makes use of eq. (17) and the government budget constraint (7). Note that $y(h)$ and $p(h)$ ($y(f)^*$ and $p(f)^*$) are the output and the price of the representative home (foreign) good. This condition simply says that the world real consumption corresponds to the world real income.

In order to analyse the effects of monetary shocks, the solution strategy is to find a flexible-price steady state as a function of relative wealth and linearise the equilibrium around it. Under a steady-state equilibrium, it is assumed that all exogenous variables are constant. These include consumption, output and the real interest rate ($\bar{r}$), which, from the consumption Euler equations (10) and (11), is given by:

$$(19) \quad \bar{r} = \delta = \frac{1-\beta}{\beta}$$

where the over-bar indicates a steady-state value and δ is the rate of time preference.

All producers in a country are also assumed to be symmetric, which implies that they set the same price and output in equilibrium. It can be shown that the steady-state per capita consumption levels are:[8]

$$(20) \quad \bar{C} = \bar{r}\bar{F} + \frac{\bar{p}(h)\bar{y}}{\bar{P}}$$

and, by making use of the identity (17), for the foreign country:

$$(21) \quad \bar{C}^* = -\bar{r}\left(\frac{n}{1-n}\right)\bar{F} + \frac{\bar{p}^*(f)\bar{y}^*}{\bar{P}^*}$$

These equations show that real consumption spending is equal to net real interest payments from abroad plus real domestic income.

In the special case of initial zero net foreign assets $\bar{F}_0 = \bar{F}_0^* = 0$, there is a closed-form solution for the steady state, in which the countries have identical per capita outputs and real money holdings. In this particular case, the first-order conditions governing each individual's optimal choice of output (15) and (16) imply that in equilibrium:[9]

$$(22) \quad \bar{y}_0 = \bar{y}_0^* = \left(\frac{\theta-1}{\theta k}\right)^{\frac{1}{2}}$$

and, from the first-order conditions (13) and (14):

$$(23) \quad \frac{\bar{M}_0}{\bar{P}_0} = \frac{\bar{M}_0^*}{\bar{P}_0^*} = \left(\frac{1-\beta}{\chi}\right)^{-1/\varepsilon} \bar{y}_0^{\,1/\varepsilon}$$

---

[8] See Appendix 1, Proof 3.
[9] See Appendix 1, Proof 4.

The first expression is equivalent to the output equation of Blanchard and Kiyotaki (1987), where producers' market power pushes global output below its competitive level (that is, with $\theta \to \infty$). Moreover, from above it can be seen that money does not matter for the steady state and, under flexible prices, output is determined independently of monetary factors.

2.1.4. The Linear Approximation Around the Symmetric Steady State

The model is now log-linearised around the initial symmetric steady state (with $\overline{F}_0 = \overline{F}_0^* = 0$), which implies expressing each variable in terms of deviations from the baseline steady-state path. Denoting percentage changes from the baseline by hats, each variable is represented by $\hat{X}_t \equiv dX_t / \overline{X}_0$, where $\overline{X}_0$ is the initial steady-state value.

The purchasing power parity relation (5) is log-linearised as:

(24) $\hat{E}_t = \hat{P}_t - \hat{P}_t^*$

where $\hat{P}_t = dP_t / \overline{P}_0$, $\hat{P}_t^* = dP_t^* / \overline{P}_0^*$ and $\hat{E}_t = dE_t / \overline{E}_0$.

Assuming symmetry among each country's producers, equations (3) and (3') yield:

$$P_t = \left\{ np_t(h)^{1-\theta} + (1-n)\left[E_t p_t^*(f)\right]^{1-\theta} \right\}^{\frac{1}{1-\theta}},$$

$$P_t^* = \left\{ n\left[\frac{p_t^*(h)}{E_t}\right]^{1-\theta} + (1-n)\left[p_t^*(f)\right]^{1-\theta} \right\}^{\frac{1}{1-\theta}}$$

Linearising around the symmetric steady state, where $\overline{p}_0(h) = \overline{E}_0 \overline{p}_0^*(f)$ the latter two expressions lead to:

(25) $\hat{P}_t = n\hat{p}_t(h) + (1-n)[\hat{E}_t + \hat{p}_t^*(f)]$

and

(26) $\hat{P}_t^* = n[\hat{p}_t(h) - \hat{E}_t] + (1-n)\hat{p}_t^*(f)$

where $\hat{p}_t(h) \equiv dp_t(h)/\overline{p}_0(h)$ and $\hat{p}_t^*(f) \equiv dp_t^*(f)/\overline{p}_0^*(f)$.

The log-linearised version of the world demand schedules for representative home and foreign products eq. (8) are:

(27) $\hat{y}_t = \theta[\hat{P}_t - \hat{p}_t(h)] + \hat{C}_t^w$

and

(28) $\hat{y}_t^* = \theta[\hat{P}_t^* - \hat{p}_t^*(f)] + \hat{C}_t^w$

Taking a population-weighted average of eqs. (27) and (28) (this implies multiplying both sides of eq. (27) by n and both sides of eq. (28) by *(1-n)* and summing the resulting equations), all the price terms cancel by eqs. (25) and (26). From the above relationships, it is easy to derive the following world goods market equilibrium condition:

(29) $\hat{C}_t^w = n\hat{C}_t + (1-n)\hat{C}_t^* = n\hat{y}_t + (1-n)\hat{y}_t^* = \hat{y}_t^w$

Equations (15), (16), (10), (11), (13) and (14) derived from the optimisation problem are approximated by:

(30) $(\theta+1)\hat{y}_t = -\theta\hat{C}_t + \hat{C}_t^w$

(31) $(\theta+1)\hat{y}_t^* = -\theta\hat{C}_t^* + \hat{C}_t^w$

(32) $\hat{C}_{t+1} = \hat{C}_t + (1-\beta)\hat{r}_t$

(33) $\hat{C}_{t+1}^* = \hat{C}_t^* + (1-\beta)\hat{r}_t$

(34) $\hat{M}_t - \hat{P}_t = \frac{1}{\varepsilon}\hat{C}_t - \frac{\beta}{\varepsilon}\left(\hat{r}_t + \frac{\hat{P}_{t+1} - \hat{P}_t}{1-\beta}\right)$

(35) $\hat{M}_t^* - \hat{P}_t^* = \frac{1}{\varepsilon}\hat{C}_t^* - \frac{\beta}{\varepsilon}\left(\hat{r}_t + \frac{\hat{P}_{t+1}^* - \hat{P}_t^*}{1-\beta}\right)$.

2.1.5 Comparing for Steady States under Flexible Prices

Before proceeding with the solution of this log-linearised model, it is important to solve for the effect of a shift in the distribution of world wealth $d\overline{F} / \overline{C}_0^w$ (since $\overline{F}_0 = 0$), in that these effects are crucial factors in the determination of the short and long-run effects of monetary disturbances. In particular, OR show that the closed-form solutions for the effects of wealth transfers on the steady state can be obtained by linearising equations (20) and (21). This leads to:

(36) $\hat{\overline{C}} = \overline{r}\frac{d\overline{F}}{\overline{C}_o^w} + \hat{\overline{p}}(h) + \hat{\overline{y}} - \hat{\overline{P}}$

and

(37) $\hat{\overline{C}}^* = -\left(\frac{n}{1-n}\right)\overline{r}\frac{d\overline{F}}{\overline{C}_o^w} + \hat{\overline{p}}^*(f) + \hat{\overline{y}}^* - \hat{\overline{P}}^*$

Note that these equations are valid only for steady-state changes and are not time subscripted. From the above equations, it is possible to see that an exogenous increase $d\overline{F}$ in home per capita foreign assets would increase steady-state consumption by the amount $\overline{r}d\overline{F}$.

Knowing that equations (25)-(35) also hold across steady states, the full system is then given by equations (25)-(37). In particular, the solutions for consumption are:[10]

(38) $\hat{\overline{C}} = \frac{1+\theta}{2\theta}\left(\frac{\overline{r}d\overline{F}}{\overline{C}_0^w}\right)$

and

(39) $$\hat{\bar{C}}^{*} = -\frac{n}{1-n}\left(\frac{1+\theta}{2\theta}\right)\left(\frac{\bar{r}d\bar{F}}{\bar{C}_0^w}\right).$$

From these equations, it is possible to see that an exogenous increase in home per capita foreign assets $d\bar{F}$ would increase steady-state consumption by less than the amount $\bar{r}d\bar{F}$ (since $\theta > 1$). This is due to the fact that higher wealth leads to some reduction in work effort and production as shown in equation (30).

As emphasised above, an important implication of the model is that the long-run equilibrium of the economy has been determined without making use of the money-demand equations (34) and (35). As a result, the classical invariance of the real economy with respect to monetary factors holds in the *Redux* model. This implies that, with prices and nominal exchange rate free to adjust immediately, monetary policy has no short-run effects on output and consumption. Variations in the nominal money stock, however, will affect the former variables proportionally.

Across steady states, inflation and interest rate do not change, so (34) and (35) become:

(40) $$\hat{\bar{P}} = \hat{\bar{M}} - \frac{1}{\varepsilon}\hat{\bar{C}}$$

and

(41) $$\hat{\bar{P}}^{*} = \hat{\bar{M}}^{*} - \frac{1}{\varepsilon}\hat{\bar{C}}^{*}$$

Subtracting eq. (41) from eq. (40) and applying the PPP eq. (24) gives the long-run nominal exchange rate:

$$\hat{\bar{E}} = (\hat{\bar{M}} - \hat{\bar{M}}^{*}) - \frac{1}{\varepsilon}(\hat{\bar{C}} - \hat{\bar{C}}^{*}).$$

[10] See Appendix 1, Proof 5.

2.2 *The Two-Country Model with Sticky Prices*

In the short-run, OR assumes that nominal producer's prices $p(h)$ and $p(f)^*$ are predetermined, that is, they are set a period in advance, but can be adjusted fully after one period. Possible sources of stickiness may be caused by menu costs of price adjustment, as emphasised by Mankiw (1985) or Blanchard and Kiyotaki (1987).

2.2.1 *Short-Run Equilibrium Conditions and Current Account Dynamics*

Two important assumptions of this model are that prices are rigid in the short-run and that producers are monopolists. With pre-set nominal prices, output becomes demand-determined for small enough shocks, because monopolists always price above marginal costs, and it is profitable to meet unexpected demand shocks at the pre-set price. In the short run, then, the equations equating marginal revenue and marginal costs in the flexible price case (30) and (31) need not hold. Instead, output is determined entirely by the demand equations (27) and (28).

Another basic implication of price inertia is that, although $p(h)$ and $p(f)^*$ are set one period in advance, the aggregate price indexes in each country will fluctuate with the nominal exchange rate. As a result, a nominal depreciation raises the domestic price index P by increasing the domestic-currency price of foreign-produced goods.[11]

With $p(h)$ and $p(f)^*$ fixed in the first period, equations (25) and (26) become:

(42) $\hat{P} = (1-n)\hat{E}$

and

(43) $\hat{P}^* = -n\hat{E}$

where hatted variables without time subscripts or overbars denote short-run (one period) deviations from the symmetric steady-state path.

Combining these price changes with (27) and (28), it is possible to show that short-run aggregate demands can be expressed as:

(44) $\hat{y} = \theta(1-n)\hat{E} + \hat{C}^{w}$

and

(45) $\hat{y}^{*} = -\theta n\hat{E} + \hat{C}^{w}$

where $\hat{C}^{w}$ is given by (29). The remaining equations of the short-run equilibrium include (32)-(35), which always hold.

Following each specific policy change, the world economy reaches its new steady state after a single period. Thus, all *(t+1)*-subscripted variables in the linearised consumption-Euler and money demand equations (32)-(35) can be replaced with steady-state changes. All *t*-subscripted variables in (32)-(35) are now interpreted as short-run values.

In the previous section it is found that the steady-state conditions (20) and (21) imply that, in the long run, income equals expenditure, so that current accounts are balanced. In the short-run, however, the home country run a current-account surplus given by:

$$CA_t = (F_t - F_{t-1}) = r_{t-1}F_{t-1} + \frac{p_t(h)y_t}{P_t} - C_t$$

Thus, since $F_0 = 0$, the linearised short-run (or period 1) current-account equations are:[12]

[11] As pointed out by Walsh (1998), this process introduces a new channel (absent in a closed economy model) through which monetary disturbances can have an immediate impact on the general price level even in the presence of nominal rigidities.

[12] See Appendix 1, Proof 6.

(46) $$\frac{d\overline{F}}{\overline{C}_o^w} = \hat{y} - \hat{C} - (1-n)\hat{E}$$

and

(47) $$\frac{d\overline{F}^*}{\overline{C}_o^w} = \hat{y}^* - \hat{C}^* + n\hat{E} = -\left(\frac{n}{1-n}\right)\frac{d\overline{F}}{\overline{C}_o^w}$$

From eq. in (46) and (47), given one-period price setting, it is easy to verify that whatever net foreign asset stock arises at the end of the first period, it becomes the new steady-state level from period 2 onwards. In formal terms, this implies that:

$$\frac{d\overline{F}_t}{\overline{C}_0^w} = \frac{d\overline{F}}{\overline{C}_0^w}, \quad \forall t \geq 2$$

The above expression represents a crucial link between the equations of the short-term and the long-term equilibrium. In fact, as shown from equations (36) and (37), the solution for steady-state variables is a function of $d\overline{F}$, which in this case is affected by short-run current-account imbalances.

2.2.2 The Effects of an Unanticipated Domestic Monetary Shock

The effects of an unanticipated permanent increase in the relative home money supply occurring in period 1 are now analysed. In particular, the full system is characterised by ten short-run variables ($\hat{C}, \hat{C}^*, \hat{y}, \hat{y}^*, \hat{P}, \hat{P}^*, \hat{E}, \hat{C}^w, \hat{r}$ and $d\overline{F}$) to be determined by the ten-equations given by (29), (32)-(35), and (42)-(46). A direct solution is possible, but, as suggested by OR, a more intuitive way is one which exploits the model's symmetry. Subtracting the foreign Euler equation (33) from its home counterpart (32) gives:

(48) $$\hat{C} - \hat{C}^* = \hat{\overline{C}} - \hat{\overline{C}}^*$$

This equation says that all shocks have permanent effects on the difference between home and foreign per capita consumption, implying the usual random walk hypothesis of consumption.

Similarly, the money demand equations (34) and (35) can be arranged as:

$$(49) \quad (\hat{M}-\hat{M}^{*})-\hat{E}=\frac{1}{\varepsilon}(\hat{C}-\hat{C}^{*})-\frac{\beta}{(1-\beta)\varepsilon}(\hat{\bar{E}}-\hat{E})$$

which states that relative money demand depends upon consumption differences, which is a common result to many other intertemporal monetary models.

An unanticipated permanent rise in the relative home money supply has immediate effects on the nominal exchange rate. This can easily be seen by taking eq. (49) and leading it by one period to obtain:

$$(50) \quad \hat{\bar{E}}=(\hat{\bar{M}}-\hat{\bar{M}}^{*})-\frac{1}{\varepsilon}(\hat{\bar{C}}-\hat{\bar{C}}^{*}),$$

which is simpler than (49) because all variables are constant in the assumed steady state. Using eq. (50) to substitute for $\hat{\bar{E}}$ in eq. (49) and eq. (48), and knowing that $\hat{M}-\hat{M}^{*}=\hat{\bar{M}}-\hat{\bar{M}}^{*}$ (since the money supply shock is permanent) leads to:

$$(51) \quad \hat{E}=(\hat{M}-\hat{M}^{*})-\frac{1}{\varepsilon}(\hat{C}-\hat{C}^{*})$$

Comparing (50) with (51) gives $\hat{E}=\hat{\bar{E}}$, implying that the nominal exchange rate jumps immediately to its new long-run equilibrium following permanent relative money shocks.

The intuition behind this result is that if consumption differentials and money differentials are both expected to be constant, then agents must expect a constant nominal exchange rate as well. If relative consumption levels do not adjust, a permanent change in $\hat{E}$ is just equal to the change in nominal money supplies ($\hat{M}-\hat{M}^{*}$).

Therefore, an increase in $\hat{M}$ relative to $\hat{M}^*$ generates a nominal depreciation of the home-country currency (a rise in $\hat{E}$). If relative consumption levels adjust, that is, $(\hat{C}-\hat{C}^*)\neq 0$, this will affect the relative demand for money given by eq. (49). As a result, equilibrium between home money supply and home money demand can be restored with a smaller increase in the home price level, which, since $p(h)$ and $p^*(f)$ are fixed for one period, generates a smaller depreciation of the exchange rate. In general, the larger the rise in home consumption, the higher is the increase in real money demand,[13] and the smaller the rise in the home price level and, then, the smaller the depreciation of the domestic nominal exchange rate.

In Figure 1.1, equation (51) is represented by the schedule *MM*, which has a slope equal to $-1/\varepsilon$. The curve is downward sloping because an increase in relative home consumption raises home money demand and home's relative price level must fall, implying an appreciation of the exchange rate. The *MM* schedule intersects the vertical $\hat{E}$ axis at $(\hat{M}-\hat{M}^*)$, which would be the equilibrium exchange rate response if prices were fully flexible. Before the monetary shock, the relevant *MM* schedule passes through the origin.

So far the relationship between $\hat{E}$ and $(\hat{C}-\hat{C}^*)$ has been determined. The real consumption differential, however, is itself endogenous. A second schedule in $\hat{E}$ and $(\hat{C}-\hat{C}^*)$ is derived by using the current-account equations (46) and (47) together with the long-run consumption equations (38) and (39). The resulting relationship leads to:

[13] Note that the real demand for money depends upon the elasticity of money demand, which is inversely related to ε. The greater ε is, the smaller the change in the money demand deriving from a change in consumption.

$$\hat{\bar{C}} - \hat{\bar{C}}^* = \frac{\bar{r}(1-\theta)}{2\theta}[(\hat{y}-\hat{y}^*) - (\hat{C}-\hat{C}^*) - \hat{E}]$$

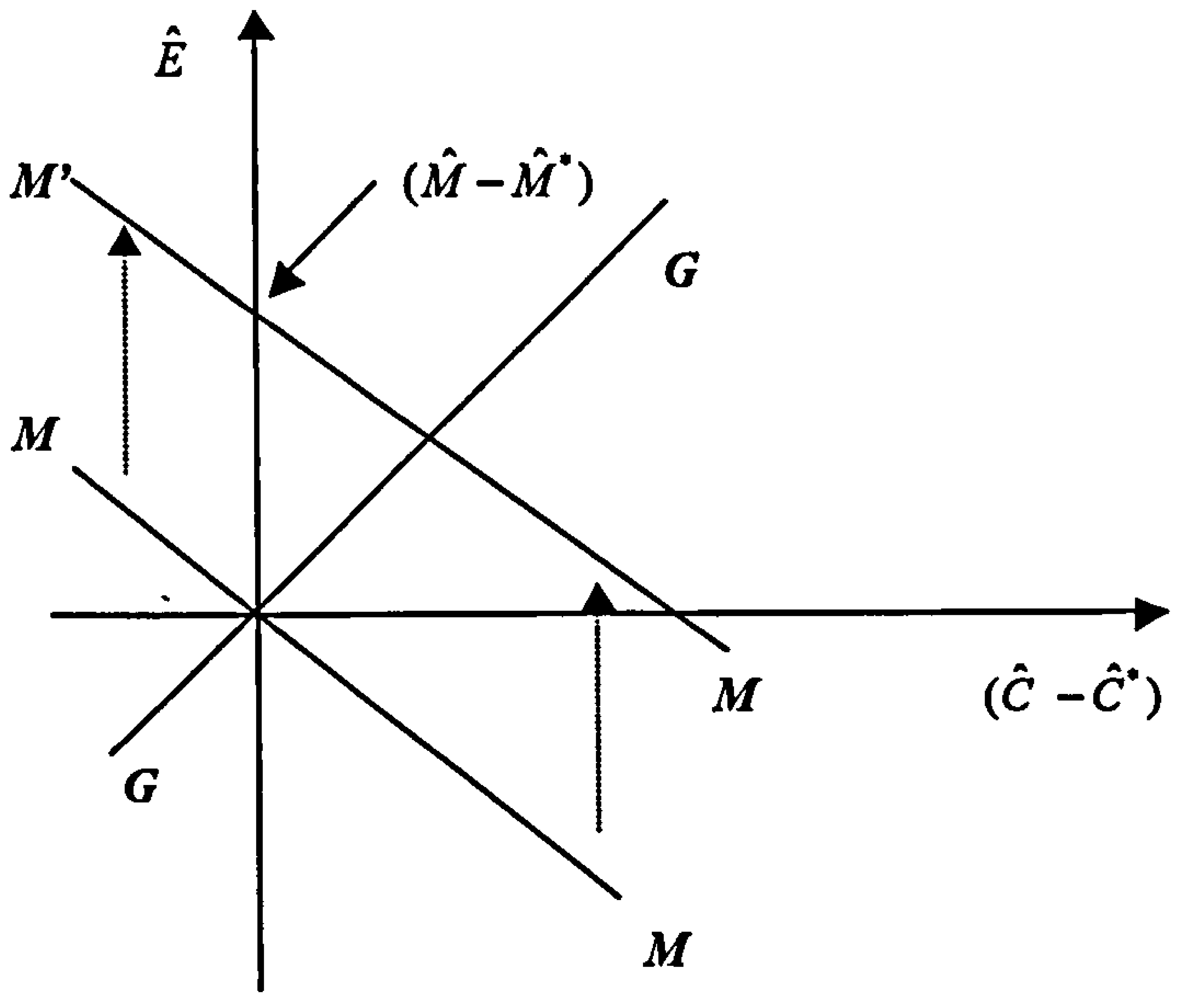
By substituting $(\hat{y}-\hat{y}^*)$ with $\theta\hat{E}$ (this is directly derived from subtracting eq. (45) and eq. (44)) and combining the resulting equation with the relative Euler equation (48)), the ***GG*** schedule is defined as:

$$\text{(52)} \qquad \hat{E} = \frac{\bar{r}(1+\theta)+2\theta}{\bar{r}(\theta^2-1)}(\hat{C}-\hat{C}^*)$$

which is upward sloping because home consumption can rise relative to foreign consumption only if the nominal exchange rate depreciates in the short run and allows home output to rise relative to foreign output. With nominal prices fixed in the short run, in fact, the initial currency depreciation switches world demand toward domestic products and causes a short-run rise in relative domestic income. This can be easily verified from the expression $(\hat{y}-\hat{y}^*) = \theta\hat{E}$, which is obtained from eq. (45) and eq. (44). Via the usual intertemporal consumption-smoothing channel, the home country residents save part of this extra income and run a current-account surplus. The higher long-run wealth and the relative higher interest income lead to substitution from work into leisure, a fall in the supply of home goods (and a rise in the relative home output prices), and a consequent permanent improvement in home's terms of trade.[14]

[14] The positive correlation between the current account and the terms of trade is put forward by Obstfeld and Rogoff in support of the 'transfer problem.' As shown by Lombardo (2002), however, this conclusion is not robust to alternative assumptions on the elasticity of substitution between domestic and imported goods. See the above reference for a full discussion.

FIGURE 1.1 Effect ofan Unexpected Money Shock on the Exchange Rate



Sources: Obstfeld and Rogoff (1996), page 680.

Figure 1.1 shows the shift of the initial pre-shock *MM* schedule to ***M'M'***, which occurs when there is a permanent unanticipated relative home money supply shock of size $(\hat{M}-\hat{M}^*)$. The intersection of ***M'M'*** and ***GG*** is the short-run equilibrium. The domestic currency depreciates, but by an amount proportionally smaller than the increase in the relative home money supply. The exchange rate change is smaller the less monopoly power producers have, or the larger the price ebsticity of demand θ. In fact, as domestic and foreign goods become close substitutes, small changes in exchange rate lead to very big shifts in demand with pre-set prices.

The equilibrium nominal exchange rate change is obtained combining eqs. (51 and (52) as:

$$(53) \quad \hat{E}=\frac{\varepsilon[\bar{r}(1+\theta)+2\theta]}{\bar{r}(\theta^2-1)+\varepsilon[\bar{r}(1+\theta)+2\theta]}(\hat{M}-\hat{M}^*)\leq(\hat{M}-\hat{M}^*)$$

whereas the relative consumption is given by:

$$(54) \quad (\hat{C}-\hat{C}^{*}) = \frac{\varepsilon\bar{r}(\theta^{2}-1)}{\bar{r}(\theta^{2}-1)+\varepsilon[\bar{r}(1+\theta)+2\theta]}(\hat{M}-\hat{M}^{*})$$

As pointed out by OR, the main implication of this model is that as long as any type of short-run nominal rigidities exists, unanticipated money shocks are likely to lead to international capital flows. The resulting transfers will extend the real effects of the shock beyond the initial sticky-price time horizon.[15] In this model, given the assumption of infinitely lived agent with intertemporally separable utility, the real effects are permanent, but in an overlapping generation version they would eventually die down (even though they will last longer than the price rigidity horizon).

To find the short-run current account effect of monetary shock, equations (38) and (39) are used to solve for $\hat{\bar{C}}-\hat{\bar{C}}^{*}$ as a function of $\frac{d\bar{F}}{\bar{C}_{o}^{w}}$; then, by taking eqs. (48) and (54), it is possible to obtain the following relationship:

$$(55) \quad \frac{d\bar{F}}{\bar{C}_{o}^{w}} = \frac{2\theta\varepsilon(1-n)(\theta-1)}{\bar{r}(\theta^{2}-1)+\varepsilon[\bar{r}(1+\theta)+2\theta]}(\hat{M}-\hat{M}^{*})$$

where it can be seen that the larger the home country (the larger n), the smaller the impact of home money supply on its current account. This is due to the effects that nominal depreciations have on the relative price level and the correspondent competitiveness gains, as shown in eq. (42) and (43). Eq. (55) summaries the main predictions of the model that will be tested in section 5.

[15] This important point is further discussed in Lee and Chinn (2002).

In order to provide a graphical representation of the overall effects of expansive monetary shocks on the current account, Figure 1.2 shows the schedule ***ZZ***, which is directly derived from eq. (55) and eq. (54), that is:

(56) $$\frac{d\overline{F}}{\overline{C}_0^w}=\frac{2\theta\,(1-n)}{\overline{r}(\theta-1)}\left(\hat{C}-\hat{C}^*\right)$$

The other important relationship comes directly from the budget constraint and is obtained by using eqs. (46) and (47). It can be written as:

(57) $$\frac{d\overline{F}}{\overline{C}_0^w}\left(\frac{1}{1-n}\right)=(\theta-1)\hat{E}-\left(\hat{C}-\hat{C}^*\right)$$

which says that the change in bond holdings (namely, in current account) is a positive function of the exchange rate deviation from the steady-state value (being $\theta > 1$) and a negative function of the difference between home and foreign per capita consumption. The first member of the RHS is a result consistent with the Marshall-Lerner condition. In addition, however, there is a consumption effect according to which the larger the increase of relative consumption, the smaller the current-account effect.[16]

Eq. (57) can be rearranged to obtain the schedule EE:

(58) $$\frac{d\overline{F}}{\overline{C}_0^w}=(1-n)(\theta-1)\hat{E}-(1-n)\left(\hat{C}-\hat{C}^*\right),$$

which implies that the smaller the home country (i.e. the smaller n), and the larger the elasticity of substitution between varieties, the larger the impact of exchange rate depreciation on the home current account.

[16] This derives from the fact that under the assumption of monopolistic competition, domestic output temporary rises. With the permanent income hypothesis, the larger the consumption effect is, (which, however, expands less that output), the smaller the home country current-account surplus.

FIGURE 1.2 Effect of an Unexpected Money Shock on the Current Account

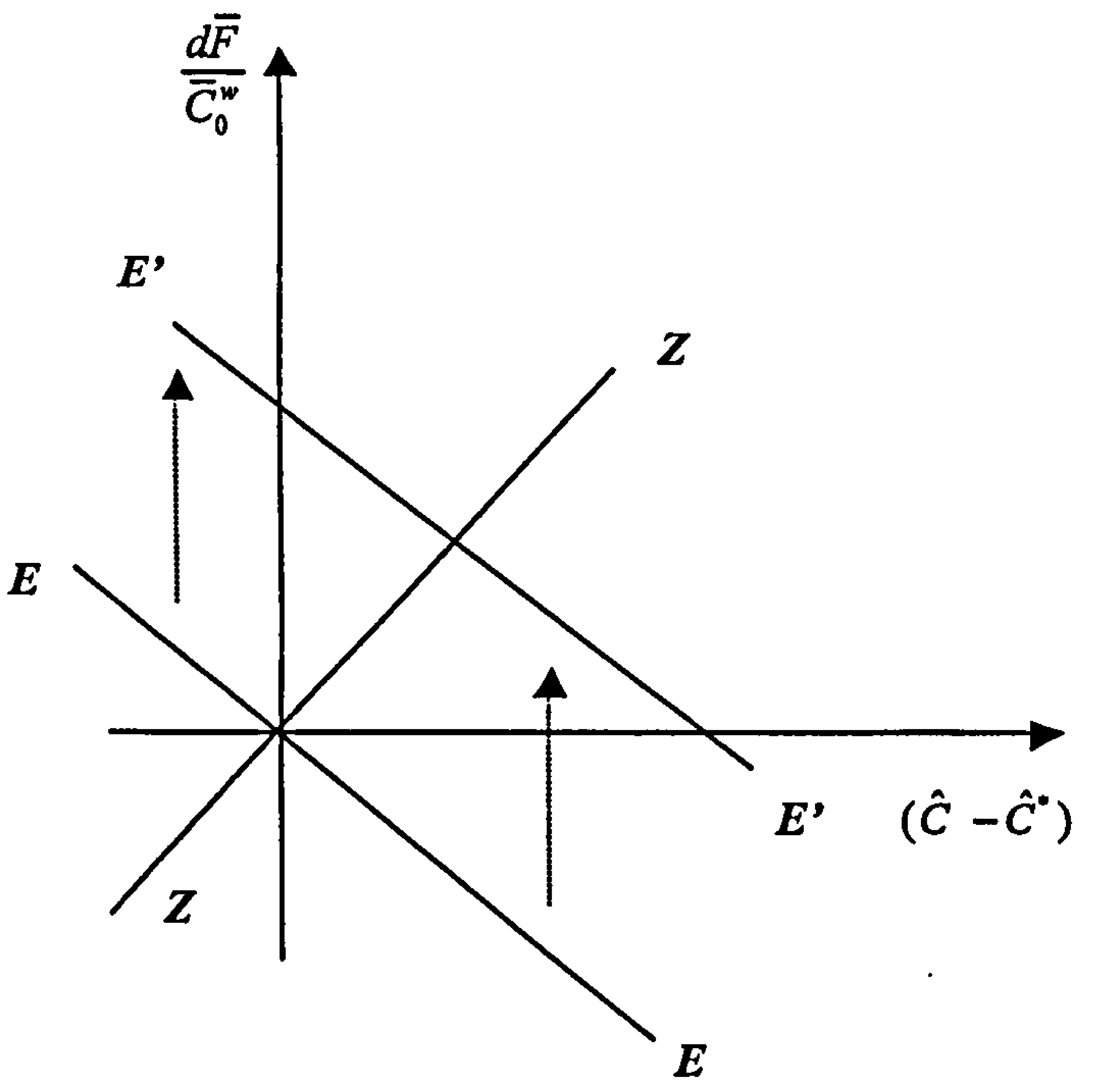


Figure 1.2 shows the effects of a home country monetary expansion. As a result of the nominal exchange rate depreciation, the schedule *EE* shifts upwards. The effects of nominal shocks on the current account are given by the intersection of the curve *E'E'* with *ZZ*.

Overall, the effects of an unexpected positive monetary policy shock can be summarised as follows. A monetary expansion in the home country produces a nominal depreciation and a subsequent rise in the domestic price level (P), followed by a decrease of the domestic relative prices ($p(h)/P$). The latter effect is due to the domestic output price rigidity. As a consequence, and under the assumption of monopolistic competition, domestic output temporarily expands. On the basis of the permanent income assumption, consumption rises less than output, leading the home country to run a current-account surplus. The excess of output over domestic consumption is exported and, as a payment for these exports, the home country receives claims against the future output of the foreign country. As a result, domestic consumption permanently rises, although the increase in output lasts for only one period.

The current-account fluctuations which follow the country-specific structural shocks constitute a key transmission channel in the *Redux* model. However, the operativeness of this mechanism does not seem robust to alternative assumptions and building hypothesis. In what follows, a brief discussion of two straightforward generalisations of the original framework is provided. These will show dramatic effects on the main predictions of the basic model, motivating the subsequent empirical sections.

3. Current Account Dynamics in Two Extensions of the Redux Model

As Obstfeld and Rogoff first pointed out, their framework could be improved through a number of extensions.[17] Lane (2001) offers an excellent survey of the recent literature attempting to extend and generalise the basic *Redux* model by introducing sticky wages, staggering nominal rigidities, market segmentation and pricing to market, different household preferences and financial structures.[18] What follows gives a brief overview of two contributions, which show how the international transmission of monetary shocks

[17] For instance, they suggest introducing overlapping generations in place of homogeneous infinitely lived agents. This assumption not only allows us to have a more realistic model than the basic one, but, more significantly, makes room for the possibility of the Ricardian Equivalence not to hold.

[18] Sarno (2000) provides another selective survey of the recent fast growing literature on new open economy macroeconomics.

and their effects on the current account and exchange rate of open economies are not robust to alternative building assumptions.[19]

Betts and Devereux (2000) extend the OR model by allowing short-run departures of the real exchange rate from PPP, which is assumed to hold in the basic framework. In particular, they introduce the hypothesis that a fraction s of firms in each country can price-discriminate across countries and, therefore, set different prices in home and foreign markets. This parameter provides a measure of price to market (PTM) in some traded goods industries, where firms tend to set prices in local currencies of sale and do not adjust prices to movements in the exchange rate.[20] In their model it is displayed that the nominal price stickiness and the presence of PTM increase the volatility of the exchange rates and strongly affect the international monetary transmission mechanisms. In particular, they formally demonstrate that a high degree of PTM reduces the traditional 'expenditure-switching' effects of exchange rate depreciation, which, as a result, has little influence on the relative price of imported goods facing domestic consumer and on the correspondent shift in world demand. The above point can be appreciated from eqs. (42) and (43) of section 2, which define the response of aggregate price indices to an exchange rate depreciation. Betts and Devereux show that in their setup the corresponding equations are given by

[19] The following generalisations of the *Redux* model are only a selected sample illustrating how the results of this class of models are sensitive to the specification of the micro-foundations. There are, however, a number of other papers in which the sign of the current account response to monetary shocks is ambiguous. For instance, Lane (1999) develops a small open economy model (extending the one outlined in the appendix of Obstfeld and Rogoff, 1995) and shows that the effects of the current account depends upon the parameters governing the willingness to substitute consumption across periods and the degree of substitutability between traded and non-traded goods. Chari, *et al.* (1998) introduce capital and show that, due to strong investment effects following reductions of the short-term interest rates, monetary shocks generate current account deficits.

[20] Following the lack of empirical support for the law of one price (LOOP) and PPP for traded goods, at least at high frequency, in the last two decades there has been a growing literature on the presence of PTM. Glodberg and Knetter (1997) provide a comprehensive survey of the empirical evidence on the degree of exchange rate pass-through, market segmentation and pricing-to-market. Recent empirical evidence is also given in Anderton (2003).

$\hat{P} = (1-n)(1-s)\hat{E}$ and $\hat{P}^{*} = -n(1-s)\hat{E}$, that is, the greater the share of goods subject to PTM, the lower the response of aggregate prices and classic expenditure-switching effect.

In particular, following a domestic monetary expansion, they find that with $s \to 0$ (e.g. the law of one price is maintained continuously and PPP holds) for both countries, a devaluation unambiguously improves the current account and generates permanent positive effects in the home consumption relative to foreign consumption. This result is the special case considered in the OR model. When $s \to 1$ and full-PTM occurs, both real and nominal exchange rates depreciate in the short run, but the current account is unchanged because domestic income and consumption rise by equal proportions. The latter effect is due to the fact that, although home individuals would like to save some of the higher income into future consumption, there is a fall in the real interest rate which pushes them to consume the full amount in the present. In general, following an unanticipated positive monetary shock, the higher the degree of PTM, the smaller the current-account effects and, as a result, the lower the permanent effects on relative consumption.

Another important extension has been developed by Tille (2001), who allows for the elasticity of substitution between two goods produced domestically to differ from the elasticity of substitution between two goods produced in different countries.[21] This generalisation is empirically relevant because countries tend to specialise in the production of some goods and it is likely that the substitutability between goods

[21] Corsetti and Pesenti (2001) also consider the special case in which the elasticity of substitution between two goods in the two different countries is equal to one. Their analysis, however, does not permit dynamic effects through the current account. This approach has been recently criticised by Cavallo and Ghironi (2002).

produced in different countries might be less that between goods produced domestically. In particular, he generalises the base line *Redux* model, where the two parameters are equal and constrained to be bigger than one (e.g. $\theta > 1$), and, under the assumption of law-of-one-price and PPP, distinguishes two possible values of the elasticity of substitution between home and foreign goods (respectively greater and less than unity). He labels the first case as the Marshall-Lerner-Robinson (MLR) condition, in that, when goods produced in different countries are close substitutes, a depreciation of the nominal exchange rate generates a current account surplus and a permanent rise of home consumption relative to foreign consumption, as in the *Redux* setup. When the goods produced in the two countries are poor substitutes and the cross-country elasticity is less than unity (e.g. the NON-MLR), the current account will be in deficit in the short run, determining a permanent fall in relative consumption. Closely related to the work of Tille (2001), Lombardo (2001) modifies the original OR specification with different degrees of elasticity of substitution between domestic and imported goods. He derives conditions allowing for positive, neutral and negative current account responses to a monetary expansion and currency depreciations, arguing that the standard Marshall-Lerner condition may not apply for specific intervals in the value of the relevant coefficients.

In summary, several models have been developed extending and generalising the *Redux* model. These papers open up to a wider range of predictions on the size and the sign of the effects of positive monetary shocks on the macroeconomic variables and, in particular, exchange rates and current accounts. As pointed out by Lane (2001), the above literature has been mainly focused on theoretical aspects. However, the actual effects of monetary shocks to exchange rates, current account and other real variables are primarily an empirical issue. While some exercises have been based on calibration methods and others on the estimation of the key parameters of the model, a number of papers have addressed this deficiency by estimating impulse response functions generated by VAR econometric techniques, where different identification solutions have been used. The following section provides a brief overview of the main studies, from which this chapter tries to improve upon.

4. An Empirical Review of VAR Studies in Open Economies

The two most influential papers investigating the monetary transmission mechanism in open economies on which several studies have built are Eichenbaum and Evans (1995) and Clarida and Galì (1994). Similarly to many works aimed at identifying the transmission mechanism of exogenous monetary policy changes, both papers make use of structural VAR models, but apply alternative identification solutions. In particular, they investigate the impact of monetary policy shocks on the real and nominal exchange rates in the G7 countries finding supporting evidence in favour of the theoretical implications of traditional Keynesian aggregative models and some of the newer sticky price intertemporal models.

Eichenbaum and Evans (1995) use U.S. monthly data over the period 1974-1990 and estimate alternative VAR models, identifying exogenous monetary shocks through innovations to the federal fund rate, the ratio of non-borrowed reserves to total reserves. By including common endogenous variables (namely, industrial production, consumer price index, short-term interest rate differentials and exchange rates), they identify nominal innovations through a Cholesky decomposition identification scheme.[22] by ordering the monetary variables before the industrial production and the consumer price

[22] Some of the traditional identification schemes (including the widely used Cholesky decomposition) are discussed in chapter 2.

level, and after the exchange rate. This corresponds to the assumption that the monetary authority sets its policy instrument with current values of the first two variables in mind, while the latter do not respond contemporaneously to movements of the monetary shock.

They estimate separate VARs with respect to Japan, Germany, the United Kingdom, France and Italy, and their results show that contractionary monetary policy shocks lead to significant and persistent appreciations of the real and nominal exchange rates. Moreover, they find that these disturbances contributed for as much as 40 per cent of the overall variability of U.S. exchange rates in the post-Bretton Woods era. Betts and Devereux (1999) extend their models to a longer sample period with slightly different specifications. Imposing short-run recursive identification restrictions, they find similar results.[23]

Lane (1999) further extends their sys tem to include the trade balance over the period 1974-1996 for the U.S. with respect to all the other G7 countries. In particular, he estimates a six-variable VAR system, imposing a recursive decomposition. The impulse response functions of the trade balance with respect to a monetary expansion show a sustained surplus after a period of about a year, although the maximum impact occurs just 3-4 years after the shock.

Sims (1992) uses similar methodologies for non-US economies, but finds that identifying monetary policy shocks in these countries produces puzzling impulse responses (namely, the 'exchange rate puzzle'). In particular, positive innovations in the short-term interest rate are associated with a statistically significant and persistent

[23] In Betts and Devereux (1996), they also focus on the CPI responses to monetary policy shocks, which they find to be quite flat and barely significant. They interpret these results in support of a high degree of PTM and, therefore, as if large movements in nominal exchange rates are not reflected in import prices.

depreciation of the domestic currency. These findings might be due to the identifying restrictions which do not allow for monetary authorities to react to contemporaneous changes in the exchange rate. More recently, Kim (2000) applies non-recursive or semi-structural VAR models, as suggested by Sims (1986) and Bernanke (1986) for three small open European countries (namely France, Italy and the U.K.) over the period 1980-1996.[24] He shows that expansive monetary policy shocks lead to a nominal depreciation of the exchange rate, which results in an improvement of the trade balance in all three countries.[25]

Another part of the VAR literature has made use of long-run identification schemes *á la* Blanchard and Quah (1989).[26] An influential contribution was made by Clarida and Galì (1994), who investigate the sources of real exchange rate movements after the collapse of Bretton Woods for the U.S. *vis-à-vis* the U.K., Germany, Japan and Canada. They estimate a parsimonious structural VAR system given by three variables (namely, relative output, real exchange rate and inflation), and assume the presence of three structural shocks to supply, demand, and money. The long-run restrictions they impose are such that nominal disturbances have no long-run impact on either output or the real exchange rate, and that the demand shock is long-run neutral for relative output. They find that in two models (namely, the ones related to Japan and Germany), nominal shocks explain a substantial amount of the variance in the dollar-DM and dollar-yen real exchange rates (41% and 35% of the variance, respectively). For the U.K and Canada, results are less supportive of a relevant role of monetary impulses. However, consistent with the VAR literature based on contemporaneous restrictions, in all countries nominal

[24] Chapter 2 implements a similar identifying approach for a wider set of European countries.

[25] His results are consistent with the dominance of the expenditure-switching effect over the income-absorption effect. These two effects move the trade balance in opposite directions. As a result, the overall movement of the trade balance is determined by the dominant effect. Similar effects are found in Prasad (1999).

shocks lead to short-run real depreciation, a temporary rise in relative U.S. output and a positive effect on the relative U.S. inflation. More recently, Rogers (1999) extends Clarida and Gali's paper and, by using several VAR specifications with long-run restrictions, estimates the contribution of various structural shocks to explaining variations in the real pound-dollar rate over 100 years of data. He finds that monetary shocks accounted for 19-60% of the variance decomposition of the real exchange rate lending empirical support to the NOEM.

More strictly related to the focus of this chapter, Lane (1999) tests for the empirical relevance of a sticky-price intertemporal model, pointing to the role of monetary disturbances for current account fluctuations. In particular, he estimates a three-variable system, where he takes the ratio of the U.S. and rest-of-the-world (RoW) output, the U.S. current account to GDP ratio and the ratio of the U.S. to RoW consumer price levels. The resulting system is assumed to be driven by a sequence of three orthogonal structural shocks (labelled supply, absorption and monetary shocks) which are identified through long-run neutrality restrictions. In particular, nominal shocks are imposed to have no long-run effects on the relative output and the current account, and the absorption shock to be long-term neutral for the relative output. The estimated impulse responses of the current account to a positive monetary shock display a short-term deterioration (which Lane interprets as a J-curve effect), followed by a significant and persistent surplus which reaches its maximum impact after 10 quarters. Quantitatively, the contribution of this shock to current account volatility constitutes about half of its variation.

[26] See the next section for an outline of this identification scheme.

Similarly to the above work, Cavallari (2001) proposes a three-variable system of the ratio of domestic to world output, the ratio of current account to output and the ratio of home to foreign short-term interest rate. By using both short-run and long-run restrictions as in Galì (1992), she assumes the presence of a supply, an absorption and a monetary shock. The latter two are restricted to have no long-run effect on the output, but, in addition, the monetary disturbance cannot contemporaneously affect the latter. Estimating this system over the period 1974-1997 for all the G7 countries, she shows that the current account response to the monetary shock differs across countries. In particular, following a monetary tightening, in the U.K., Italy, France and Canada, the current account goes into surplus. For the US, Germany and Japan the reverse holds.

Although very relevant, the latter two studies present some drawbacks, which deserve some discussion. Lane (1999), for instance, justifies the long-run neutrality restriction of monetary shocks on the current account on the basis of a small-country sticky-price intertemporal model. This identifying assumption, however, provides no discriminatory power, in that, it is easy to show that in models where the stock of net foreign assets is constant in a steady state, both real and monetary disturbances have no long-run effect on the current account (Lee and Chinn, 2002). The latter identifying restriction, therefore, is not justifiable and does not help to discriminate across different structural shocks. Cavallari (2001) partly overcomes these problems by leaving the short and long-run dynamics of current account to be freely determined. It is hard, however, to disentangle demand and nominal shocks simply on the basis of a contemporaneous zero restriction of the latter on the output. This is also inconsistent with most of the intertemporal models where contemporaneous current account changes derive from a contemporaneous, but temporary, effect of monetary disturbances on the output level. Moreover, both Lane's and Cavallari's empirical models do not include any exchange rate, which is a fundamental driving variable in the international transmission mechanism of monetary disturbances to current imbalances.

In this respect, an interesting approach has been suggested by Lee and Chinn (1998, 2002) (henceforth LC), who estimate a very parsimonious two-variable VAR model containing the first difference of the real exchange rate and the ratio of the current account to GDP for the G-7 countries over the period 1980-1999. They impose a long-run neutrality restriction of a temporary (monetary) shock on the real exchange rate and show that expansive monetary policy actions determine a statistically significant temporary real exchange rate depreciation and a short-run current account surplus.

Although not directly comparable in that the focus is on the trade balance rather than the current account, Prasad (1999) provides similar conclusions. In particular, he estimates a three-variable system with the relative output level, the real exchange rate and the ratio of trade balance to output for the G-7 countries. In accordance with the implications of a modified version of Clarida and Gali (1994) model, he assumes that monetary shocks have no long-run impact on the real exchange rate and on the relative output, and that the demand shock does not affect the relative output in the long run. His results show that nominal shocks appear to have played an important role in the dynamics of the trade balance over the period 1974-1996 in most of the G-7 countries. In particular, he finds that positive monetary shocks determine significant and permanent trade surpluses.

More recently, Fisher and Huh (2002) extends Prasad's methodology without imposing any long-run PPP to identify nominal shocks. They argue that under the open economy intertemporal models with sticky prices, nominal shocks can have long-run effects on both the real exchange rate and the trade balance. Using data over the post-Bretton Woods period, for each of the G-7 countries they estimate the same three-variable system proposed by Prasad (1999), but imposing the same mix of long-run and contemporaneous restrictions as in Cavallari (2001). Their results show that nominal shocks have a positive and significant long-run effect on each country's trade balance and generate a real depreciation of the exchange rate, which in all the economies but the U.S. and Japan is permanent.[27]

Closely following the previous empirical literature, this chapter implements and estimates two structural VARs as in Blanchard and Quah (1989) and Clarida and Galì (1994) for a wide set of open economies. This approach is adopted for two main reasons. Firstly, the main interest is to assess the role of monetary disturbances on the current account with a minimum of identifying restrictions, which both the theoretical and the empirical literature seem to agree on. By applying contemporaneous identifying restrictions, the short-run dynamics of the variables in the system are inevitably constrained. Given that the latter are the main focus of this section, a more flexible approach is preferred. Secondly, this chapter seeks to test for the robustness of the results obtained by Lee and Chinn (1998, 2002), extending it to an additional set of open-economy countries (where current-account fluctuations should play a significant role in the international monetary transmission mechanism) and using a new specification. The insights and advantages of the approach followed in this chapter will become clearer in the sections below.

[27] Two points are worth pointing out. First, they find that the level of the trade balance over output ratio is stationary in some countries. Accordingly, they test the robustness of their results using both the level and the first difference, finding that in the former case the responses of the trade balances tend to dampen out at long-term horizons. Secondly, by applying one-standard error confidence intervals, it is relatively easy for them to reject the null-hypothesis that the impulse responses equal zero.

5. *Empirical Results*

5.1 Data

This chapter examines the dynamics of the current account for 15 industrialised countries using quarterly data over the pre-EMU period 1979:2-1998:4.[28] In the empirical models below, the following variables are used: the CPI-deflated real exchange rate (RER), the current account to output ratio (CA/Y) and the relative output (Y/Y*). The sources of the data are the International Monetary Fund's International Financial Statistics and the OECD Analytical Database. The definitions and construction methods are discussed in Appendix 2.

In order to understand the importance of disentangling the role of different shocks in affecting the real exchange rate and the current account fluctuations, unconditional correlations between the two variables have been calculated. In most cases, the contemporaneous correlation coefficient ranges between a low -0.05 in Austria and a relatively high -0.62 in Finland.[29] The only exceptions are Denmark, Japan, Belgium and the United Kingdom, which display contemporaneous positive correlations. In the former two countries, however, cross correlations turn negative after four and five quarters respectively. This implies that real exchange rate depreciations are generally associated with contemporaneous or lagged current account surpluses in the vast majority of the economies under investigation. While these patterns are consistent with a role of monetary shocks leading to external improvements through exchange rate effects, it is however necessary to correctly evaluate their specific role relatively to other real shocks. The analysis below introduces the structural vector

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[28] They are Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Portugal, Sp ain, Sweden, the U.K. and the U.S.

[29] The real exchange rate is defined as units of foreign currency in terms of national currency. As a result, a decrease implies a real depreciation.

autoregressive (VAR) methodology and the identification approach which will be applied to answer this question.

5.2 *The Econometric Methodology*

The empirical strategy implemented in this chapter builds on the work by Blanchard and Quah (1989) and Clarida and Galì (1994), who propose an identification scheme for VARs, in which long-run restrictions on the behaviour of the variables in the system are imposed.

In outlining this approach, a three-variable system is considered. Letting $\mathbf{x}_t \equiv [y_t, z_t, q_t]'$ denote the *(3x1)* vector of endogenous stationary variables and $\mathbf{e}_t \equiv [s_t, d_t, m_t]'$ denote the *(3x1)* vector of the system's structural shocks,[30] $\mathbf{x}_t$ can be assumed to be a covariance stationary vector process, generated by the following structural moving average (MA) model:

(59) $\mathbf{x}_t = \mathbf{C}(L)\mathbf{e}_t = \mathbf{C}_0\mathbf{e}_t + \mathbf{C}_1\mathbf{e}_{t-1} + \mathbf{C}_2\mathbf{e}_{t-2} + \ldots$

where $\mathbf{C}_0$ is the *(3x3)* matrix of the contemporaneous structural relationship among y_t, z_t and q_t, and where it is assumed that the structural disturbances $\mathbf{e}_t$ are mutually orthogonal and have unit variance, implying that $E(\mathbf{e}_t\mathbf{e}_t') = \mathbf{I}$.

The reduced-form MA representation for $\mathbf{x}_t$ is given by:

(60) $\mathbf{x}_t = \mathbf{R}(L)\mathbf{u}_t = \mathbf{u}_t + \mathbf{R}_1\mathbf{u}_{t-1} + \mathbf{R}_2\mathbf{u}_{t-2} + \ldots$

where $\mathbf{u}_t$ is a *(3x1)* vector reduced-form disturbance.

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[30] It is important to point out that, in contrast with standard identification solutions based on contemporaneous restrictions, this approach does not directly associate the three structural shocks with the three sequences {y(t)}, {z(t)} and {q(t)}. In particular, the latter can be thought of as the endogenous variables, and the three shocks as the exogenous variables of the system.

Assuming that there exists a non-singular matrix $\mathbf{S}$ such that:

(61) $\mathbf{u}_t = \mathbf{S}\mathbf{e}_t$

and comparing (59) and (60), it can be verified that:

(61.i) $\mathbf{C}(L) = \mathbf{R}(L)\mathbf{S}$ or $\mathbf{C}_0 = \mathbf{S},\ \mathbf{C}_1 = \mathbf{R}_1\mathbf{S},\ \mathbf{C}_2 = \mathbf{R}_2\mathbf{S}\ldots\ldots$

This implies that eq. (61) can be also written as:

(62) $\mathbf{u}_t = \mathbf{C}_0\mathbf{e}_t$.

Ordinary Least Square (OLS) estimation can be used to obtain consistent estimates of the parameters in (60) as well as an estimate of the symmetric variance-covariance matrix of the reduced-form disturbances $\mathbf{u}_t$:

(63) $E(\mathbf{u}_t\mathbf{u}_t') = \mathbf{S}$

Therefore, from (61), (62) and (63), and the assumption of mutually orthogonal structural shocks, together with the normalisation condition above, it can be shown that:

(64) $\mathbf{S} = \mathbf{C}_0\mathbf{C}_0' = \mathbf{S}\mathbf{S}'$.

As it is well known in the literature, the system (64) provides nine equations in only six unknowns, that is, the three variances and the three covariances that define $\mathbf{S}$. Just-identification of model (60) and, therefore, estimation of the matrix $\mathbf{C}_0$ and the structural innovations $\mathbf{e}_t$ require three additional restrictions.

In particular, letting $\mathbf{C}(1) \equiv \mathbf{C}_0 + \mathbf{C}_1 + \mathbf{C}_2 + \ldots$ denote the matrix of long-run coefficients such that:

(65) $$\mathbf{C}(1) = \begin{bmatrix} C_{11}(1) & C_{12}(1) & C_{13}(1) \\ C_{21}(1) & C_{22}(1) & C_{23}(1) \\ C_{31}(1) & C_{32}(1) & C_{33}(1) \end{bmatrix}$$

Clarida and Galì (1994) assume three long-run neutrality conditions such that $\mathbf{C}(1)$ is restricted to be lower triangular. This means that the structural shocks δ_t and v_t do not affect the variable y_t in the long run, implying that

(66) $C_{12}(1) = C_{13}(1) = 0$.

Similarly, the assumption that structural shock v_t does not influence the second endogenous variable z_t in the bng run requires that:

(67) $C_{23}(1) = 0$.

On the basis of the above set of restrictions, the matrix of long-run coefficients is now:

(68) $$\mathbf{C}(1) = \begin{bmatrix} C_{11}(1) & 0 & 0 \\ C_{21}(1) & C_{22}(1) & 0 \\ C_{31}(1) & C_{32}(1) & C_{33}(1) \end{bmatrix}$$

Moreover, it is possible to show that the restrictions given in (66) and (67) are sufficient to identify the matrix $\mathbf{C}_0$ and recover the structural-system dynamics defined by $\mathbf{C}_1, \mathbf{C}_2, \ldots$, as well as the structural shocks $\mathbf{e}_t$.

From (61.i) it easy to see that $\mathbf{R}_0 = \mathbf{I}$, $\mathbf{R}_1 = \mathbf{C}_1\mathbf{C}_0^{-1}$, $\mathbf{R}_2 = \mathbf{C}_2\mathbf{C}_0^{-1}$, and so on. Therefore, the reduced-form MA model (60) can be rewritten as:

(69) $\mathbf{x}_t = \mathbf{R}_0\mathbf{u}_t + \mathbf{R}_1\mathbf{u}_{t-1} + \mathbf{R}_2\mathbf{u}_{t-2} + \ldots$

where:

(70) $\mathbf{R}(1) \equiv \mathbf{R}_0 + \mathbf{R}_1 + \mathbf{R}_2 + .. = \mathbf{C}(1)\mathbf{C}_0^{-1}$.

Under these assumptions, Clarida and Galì (1994) define the matrix:

(71) $\mathbf{R}(1)\mathbf{S}\mathbf{R}(1)'$

which can be computed from the estimation of the variance-covariance matrix of the reduced-form disturbances $\mathbf{S}$ and the reduced-form long-run coefficients $\mathbf{R}(1)$.

Moreover, using the definition of $\mathbf{R}(1)$ and equation (61) to substitute for $\mathbf{S}$, eq. (71) can be written as:

(72) $\mathbf{R}(1)\mathbf{S}\mathbf{R}(1)' = \mathbf{C}(1)\mathbf{C}(1)'$.

Letting $\mathbf{H}$ denote the lower triangular Cholesky decomposition of equation (72) and knowing that this matrix is unique, it is possible to show that $\mathbf{H}\mathbf{H}' = \mathbf{R}(1)\mathbf{S}\mathbf{R}(1)' = \mathbf{C}(1)\mathbf{C}(1)'$, implying:

(73) $\mathbf{C}(1) = \mathbf{H}$

The latter results follow from the Cholesky decomposition of the structural long-run coefficient matrix $\mathbf{C}(1)$.

Finally, substituting (73) into (70), the following equation can be obtained:

(74) $\mathbf{C}_0 = \mathbf{R}(1)^{-1}\mathbf{H}$

which allows the computation of the structural dynamics defined by $\mathbf{C}_1$, $\mathbf{C}_2$,... as well as the structural shocks $\mathbf{e}_t$, which can be used for impulse response and variance decomposition analyses.

5.3 *Estimation Results*

The specifications and the identifying restrictions used in this chapter closely follow and extend the bivariate VAR system of Lee and Chinn (1998, 2002) and the trivariate model suggested by Clarida and Gali (1994) and Prasad (1999). In particular, in the above papers the authors identify monetary (or nominal) shocks by imposing a long-run neutrality restriction of this disturbance on the real exchange rate. Lane (1999, 2001) argues that their identification assumption is not "warranted" in the study of intertemporal sticky-price models. In particular, he points out that if monetary shocks generate current account imbalances, they will also have long-run effects on the real exchange rate.[31]

While this is true in many theoretical extensions of the *Redux* framework (i.e. models with traded and non-traded goods), as emphasised by Obstfeld and Rogoff themselves, the long-run non-neutrality should not be overstated. Firstly, as Lee and Chinn (2002) formally prove, such effects are quantitatively very small. Secondly, by allowing for overlapping generations of finite-horizon consumers, the long-run real exchange rate effects of monetary shocks dissipates. Finally, this neutrality restriction is also trivially consistent with the monetary-approach and all Keynesian aggregative models,[32] as well as the OR model itself where there is no deviation from the CPI-based purchasing power parity both in the short and long run.[33] *A priory*, therefore, one would expect monetary disturbances to have no-long-run impact on the real exchange rate.

An important advantage of this identification assumption relies on the fact that no long-run or short-run restrictions are imposed on the current account dynamics. As a result, in the multivariate models below short-term external imbalances are allowed to be freely determined. In the long run, on the other hand, one would expect any structural shock to have only temporary (although potentially persistent) effects on the current

[31] There is a large empirical literature emphasising the long-term implications of cumulated current account imbalances (or net foreign assets) on the real exchange rate (see a recent work from Lane and Milesi-Ferretti, 2000). This chapter mainly focuses on the short-run dynamics of real exchange rates and current account, and, consequently, abstracts from these long-run effects. Moreover, from a theoretical point of view, as emphasised by Obstfeld and Rogoff (1996), the long-run non-neutrality of monetary shocks would disappear in an overlapping generations (OLG) model.
[32] In the Mundell-Flemming model positive nominal shocks lowers the home interest rate and leads to a depreciation of the real exchange rate and a short-run increase in output. In the long run, however, both the real exchange rate and the output return to their initial values.
[33] In the basic OR model, free trade implies that the LOOP holds for individual goods. Moreover, given that consumers have identical preferences across goods, goods freely traded and prices set in the currency of the seller, PPP holds all times. As a result, the real exchange rate defined in terms of consumer prices is always constant both in the short and in the long run. On the other hand, the terms of trade (or the real exchange rate defined in terms of output prices) varies and represents a central equilibrating mechanism in the basic framework. On the basis of these theoretical assumptions, although containing a non-tradable component, the CPI-deflated exchange rate is taken as a proxy for the real exchange rate.

account. This is due to the fact that the unit root tests strongly reject the non-stationarity hypothesis of the current account to GDP ratio (see Appendix 2), a result which by itself is consistent with many intertemporal models.

5.3.1 Estimation of the Bivariate VAR System

The first specification is an application of the Lee-Chinn bivariate system given by the vector of endogenous stationary variables $\mathbf{x}_t = [\Delta RER_t, CA_t / Y_t]'$, which is extended to the 15 OECD countries. Their two-variable model assumes the existence of two structural shocks generating current account and real exchange rate dynamics: a permanent (p_t) and a temporary shock (t_t). In particular, the vector of stationary endogenous variables, the vector of structural disturbances, and the matrix of long-run parameters are given by:

$$(75) \qquad \mathbf{x}_t = \begin{bmatrix} \Delta RER_t \\ CA_t / Y_t \end{bmatrix} \mathbf{e}_t = \begin{bmatrix} p_t \\ t_t \end{bmatrix} \mathbf{C}(1) = \begin{bmatrix} C_{11}(1) & 0 \\ C_{21}(1) & C_{22}(1) \end{bmatrix}$$

where it is assumed that the cumulative effect of the temporary shock t_t on the growth of the real exchange rate is zero, implying that t_t has no long-run effect on the level of the real exchange rate.[34] As a result, the temporary shock t_t can be interpreted as a monetary disturbance.[35] The permanent shock p_t, on the other hand, is thought of as a real shock to productivity or preferences.

The advantage of such a model is in its simplicity, as well as in the fact that it is potentially capable of identifying the impact of two fundamental disturbances without

[34] This restriction is trivially consistent with the PPP assumption of the basic OR model and the generalisation by Tille (2001), Lombardo (2001, 2002), Betts and Devereux (2000), Chari, *et al.* (1998), where short-run fluctuations of the real exchange rate are allowed in the short run.

constraining the short-run dynamics of the current account, which are the primary focus of this chapter.

The above system is estimated independently for each economy, by taking the first difference of the real effective exchange rate (*RER*) series and the level of the ratio of the current account over output (*CA/Y*), in accordance with stationarity properties of the series. Appendix 2 shows the unit root analysis. In most of the countries, the SVARs are estimated with two lags, which appear to be a fair balance between the lag length selected by Schwartz Information Criterion (SIC) and Akaike Information Criterion (AIC). In particular, while the SIC tends to prefer the most parsimonious specification (i.e. 1-2 lags), the AIC generally suggests 2-3 or, in some cases, more lags. In order to obtain white noise residuals, four lags were necessary in Denmark and Belgium, and three in Sweden, Spain and the Netherlands. For all countries results are qualitatively robust to alternative lag lengths.

Given that the estimated impulse response functions (IRFs) of the real exchange rate and the ratio of the current account to output in response to the two structural shocks are very similar across countries, in order to save space, Figures A.3.1 (Panel A) in Appendix 3 shows the full results for a representative country, namely the U.S.. The solid line displays the impulse response, with dotted lines tracing the two-standard-deviation bands obtained with bootstrapping techniques with 1,000 replications. The responses of the level of the real exchange rate to the permanent and temporary shocks can be seen more clearly in Panel B in the same figure, in which the first-difference

[35] This specification does not allow for a distinction between money demand and money supply shocks. In order to do so, it would be necessary to include monetary instruments and implement alternative identification restrictions. See Gali (1992) for an application to U.S. data.

responses have been accumulated. A summary of all the IRFs for all the OECD-15 can be found in Table A.3.1 of Appendix 3.

The results indicate that a positive temporary shock generates a statistically significant short-run improvement of the current account to output ratio in all the countries under investigation (Figure A.3.2). The level of the real exchange rate immediately depreciates in the short term, then gradually dies out in the long run, deteriorating the current account. The latter effect, however, is not always statistically significant. As shown in Table A.3.1, the U.K. and Belgium provide some anomalous results in that the temporary current account surplus is accompanied by an appreciation of the real exchange rate. The latter response, however, is not statistically significant. In general, these patterns are consistent with the interpretation of temporary shocks as nominal shocks.

The effects of permanent disturbances are much less clear-cut. In particular, in all countries but Belgium, Denmark and the U.K., these shocks generate a temporary surplus of the current account which is associated with a real permanent appreciation of the exchange rate.[36] While the latter effect is significant at 5% level in all 15 OECD countries, however, with the exception of the U.S., Sweden, Spain, Japan, Germany and Finland the external imbalance response is not statistically different from zero. As pointed out in Lee and Chinn (2002) and on the basis of the unconditional correlations found in section 5.1, the fact that the two identified shocks generate opposite correlations between the current account and the real exchange rate indicate that that it is important to disentangle the effects of specific shocks on the two variables. At the

[36] The fact that, as the real exchange rate appreciates, for some countries a statistically significant short-run current account surplus is found might be due to the fact that the permanent shock might pick up a potential long-run effect of the monetary shock on the real exchange rate. This miss-aggregation problem is directly discussed in the following section.

same time, the mixed responses of the two variables to the permanent shock might be due to the parsimony of the system, which does not include other key variables of the transmission mechanism or distinguish between other real shocks. While supply shocks are generally thought to lead to deficits of the current account, real disturbances could also include independent 'demand' or 'preference' shocks (towards home exports) which could lead to a current account surplus. The presence of shocks affecting the variables in different directions has been subject to several criticisms of the long-run restrictions VARs (Faust and Leeper, 1997).[37] In order to test for the robustness of the above results, the next subsection implements a less parsimonious specification, which aims to improve the identification of the 'fundamental' disturbances.

Figure A.3.3 shows the computed forecast error variance decompositions (FEVD) of the current account. In most cases nominal shocks play a greater role in explaining the variation of the current account during the first few quarters after the shock. In particular, with the only exception of the U.S. (where the role of monetary disturbances for the current account is around 10-20%) and Spain, Sweden and Finland with variance decompositions progressively increasing from 35-45% during the first year to around 60% in the longer term, this shock accounts for more than 80% of the current account volatility in all the other OECD economies. Contemporaneously (results not shown), permanent shocks have a dominant role in explaining the real exchange rate variance in all countries.

[37] In the Blanchard and Quah methodology if one identified the permanent shock as two or more aggregate independent shocks which affect the underlying macroeconomic variables in different directions, then the identification of the former shock is not valid. See Rogers (1999) for a discussion.

5.3.2 Estimation of the Trivariate VAR System

In order to verify the robustness of the above results and provide greater comparability with the Clarida-Galì and Prasad studies, the above bivariate specification is augmented with a key endogenous variable entering the transmission mechanism, that is, the relative output Y/Y^*, which is defined as the ratio of domestic to world (OECD) output.[38] The inclusion of the relative output in this specification is also motivated by two main reasons. Firstly, in the theoretical models of section 2 and 3, the effects of nominal disturbances on the real activity of the home country are discussed relative to the foreign country, which in the current analysis is represented by the rest of the world. Secondly, the current account and the real exchange rate fluctuations in open economies can be thought of being largely dependent on external factors.

According to the unit root tests, the three stationary endogenous variables forming the new system are the first differences of the (logarithms of) relative output and the real effective exchange rate, and the level of the ratio of the current account to GDP. As a result, the vector of endogenous variables is given by $\mathbf{x}_t = \left[\Delta(Y_t / Y_t^*), \Delta RER_t, CA_t / Y_t\right]'$.

The above specification is chosen for a number of reasons. First, as pointed out above, the Lee-Chinn two-variable system could be subject to problems of miss-aggregation of multiple shocks (Faust and Leeper, 1997). This might be suggested by the fact that the impulse responses to permanent shocks tend to generate some mixed results. In this regard, a trivariate model allows for the identification of an additional real shock, namely a demand or absorption disturbance which is assumed to have only temporary effects on the relative output as in Clarida and Galì (1994), Lane (1999) and

[38] See Appendix 2 for a definition.

Prasad (1999).[39] In doing so, not only an additional driving disturbance is introduced, but it could 'improve' the correct identification of the monetary shock, which, consistently with a number of empirical and theoretical models has no long-run effect on the real exchange rate and output. Namely, the vector of stationary endogenous variables, the vector of structural disturbances, and the matrix of long-run parameters are given by:

$$(76) \quad \mathbf{x}_t = \begin{bmatrix} \Delta(Y_t / Y_t^*) \\ \Delta RER_t \\ CA_t / Y_t \end{bmatrix} \mathbf{e}_t = \begin{bmatrix} s_t \\ d_t \\ m_t \end{bmatrix} \mathbf{C}(1) = \begin{bmatrix} C_{11}(1) & 0 & 0 \\ C_{21}(1) & C_{22}(1) & 0 \\ C_{31}(1) & C_{32}(1) & C_{33}(1) \end{bmatrix},$$

where s_t, d_t and m_t are the relative supply, demand and monetary structural shocks respectively.

The system is estimated individually for each of the 15 OECD countries using a number of lags, which represent a good balance between the selection made by the AIC and the SIC. Accordingly, a VAR(4) is chosen for Sweden, a VAR(3) for Portugal and Germany and a VAR(2) for the remaining countries. The main results are, however, qualitatively robust to alternative lag lengths.

A representative illustration (e.g. the United States) of the full system dynamics to the three fundamental impulses is given in Figure A.3.4 of Appendix 3, whereas a descriptive overview of the IRFs in each of the OECD-15 countries is reported in Table A.3.2. The identified permanent shock, which could be interpreted as a positive supply or technology shock, permanently and significantly increases the level of relative output in all countries. The same disturbance generates permanent effects on the real exchange rate, but the sign differs across the sample. In particular, seven countries (namely,

[39] The demand shock is identified in a similar manner also by Rogers (1999), who, however, uses a much less parsimonious specification with five endogenous variables and ten long-run identifying restrictions.

Austria, Belgium, Denmark, the Netherlands, Japan, Sweden and the United Kingdom) display a real depreciation, which is statistically significant only in the latter three. A permanent and generally significant appreciation is found in the remaining economies. Mixed effects of supply disturbances on the real exchange rate are also reported in Prasad (1999) and Clarida and Galì (1994). From a theoretical point of view, technology improvements in an aggregative MFD model are expected to produce permanent real depreciations. On the other hand, both intertemporal models and real exchange rate determination based on the Balassa-Samuelson framework predict real appreciations. The estimated IRFs clearly indicate that it is not possible to provide a clear-cut answer to which model is empirically more relevant.[40]

The responses to the identified demand shocks are clearer. In particular, it is found that they generate a temporary (but only if a few case statistically significant) improvement of the level of the relative output and a permanent real appreciation of the exchange rate in all countries. Although using different specifications and identifying restrictions, the same results are also found in Prasad (1999), Clarida and Galì (1994) and Fisher and Huh (2002). In most OECD-15, the nominal shocks lead to short-run increases in the level of the relative output, which is, however, statistically different from zero in very few countries.[41] As in the bivariate specification, in all models but the

[40] Muscatelli, *et al.* (2001) present empirical evidence on the forces driving real exchange rates by using data over 120 years for three industrialised countries and, applying a time-varying VAR approach, find mixed results. They justify their findings arguing that different responses might be consistent with the alternative prevailing theoretical model.

[41] The lack of significant output effects in the three-variable VAR could imply a weak operativeness of the current account channel. There are, however, two possible explanations for these empirical findings. The first is that the relative output has been constructed as the ratio of domestic output over the OECD output. For relatively large countries (for instance the U.S., Japan and Germany) it is evident that the effect on domestic output of the structural shocks is biased downwards because in its measure (Y/Y^*) the denominator contains also the national output. This point is explained in the Data Appendix. An additional reason is that the current account normalised with the domestic output. As a result, a substantial part of the effects of structural shocks on output is implicitly accounted for in this variable, reducing the overall statistical significance of the relative output.

United Kingdom and Belgium expansive monetary shocks are associated with significant short-run real depreciations.

In order to enhance comparability with the two-variable system, Figure A.3.5 displays the dynamics of the current account to the three structural shocks in all 15 countries. The newly estimated impulse responses to the nominal disturbances produce results which are remarkably similar to the ones of the bivariate specification displayed in Figure A.3.2. In particular, positive monetary impulses cause temporary, but statistically significant improvements of the current account in all countries. Additionally, with the only exception of Portugal, the size of the responses between the two models is very similar. This seems to suggest that the two-variable system is capable of properly identifying the nominal disturbances.

As for the effects of the supply and demand shocks, the findings are more mixed. With only few exceptions (e.g. Denmark, Finland, Germany and Spain), supply disturbances generate temporary deficits of the current account, whereas demand shocks lead to small surpluses. Although these responses are statistically significant in very few circumstances, it is evident that the results related to real disturbances are not clear-cut. This might suggest the need to estimate alternative models where other independent real shocks could be more clearly disentangled and identified. For instance, it might be worth further extending this approach with the inclusion of other variables as in Rogers (1999). Under the long-run restriction methodology, however, it becomes difficult to justify additional neutrality restrictions. It is, however, very encouraging that the IRFs of the nominal shocks remain unaffected in the three-variable system, providing supportive evidence in favour of a correct identification of these disturbances.

Figure A.3.6 shows the variance decomposition of the current account. With the exception of Finland, Germany and the U.S. (around 10-20%), and Sweden, Denmark and Japan (circa 40%), monetary disturbances account for more than 50-60% of the variance of the current account over most of the 24-quarter-period horizon.[42] Similarly to the bivariate model, real shocks tend to explain most of the variance of the real exchange rate. Nominal disturbances, however, play a significant part in the FEVD of the exchange rate in many countries: 80% in Sweden, 40-50% in Italy and Spain, 20-25% in the U.K, Portugal and Finland, and 5-10% in the remaining economies.

Although consistent with the previous studies, these findings could be further tested with alternative approaches better suited to disentangle money supply from money demand shocks. As pointe d in Clarida and Gali (1994), a nominal shock that causes a temporary real depreciation and rise of the relative output is a feasible outcome of both increases of the domestic money supply and reductions of the domestic money demand relative to the rest of the world. As a result, the identified nominal shock might aggregate the two shocks all together. Lane (1999) argues that, under the long-run restriction approach, shocks are identified by the pattern of macroeconomic outcomes, which implies a wider definition of a monetary shock. An alternative methodology would be to estimate standard unrestricted VAR models in which monetary impulses are identified with exogenous changes of a specific policy instrument (e.g. short-term interest rates). As discussed in the empirical review section, however, this would imply the imposition of contemporaneous restrictions among the endogenous variables. Given

[42] For the U.S., Lane (1999) shows that the identified monetary shock accounts for about 50% of the variance of the current account. Similarly, Prasad (1999) and Fisher and Huh (2002) report that in the G-7 economies nominal disturbances dominate the variance decomposition of trade balance in most cases. The results of this chapter are also consistent with Cavallari (2001) who finds that monetary shocks drive more than half of the current account variation Italy, the U.K., France and Germany, and Canada, while accounting for less than 30% of total variance in Japan and the U.S.

that the focus of this chapter is on the short-run dynamics of the current account, the long-run identification seems a more adequate approach.

In general, the above findings are consistent with those models of the 'New Open Economy Macroeconomics' literature where the current account imbalances represent a key transmission channel of monetary impulses. Therefore, in modelling open economies ignoring these effects could constitute an important miss-representation. Additionally, these results strongly support the basic prediction of the *Redux* model and those extensions characterised high elasticity of substitution between domestic and imported goods, in which expansionary monetary shocks unambiguously generate current account surpluses. At the same time, however, it is shown that monetary shocks produce statistically significant real depreciations of the exchange rate, which is in line with the presence of some degree of price to market behaviour as predicted in many NOEM models.

5.3.3 Degree of Openness and Current Account Effects

This final sub-section focuses on an additional prediction of the *Redux* model, which refers to the relationship between country size and current account effects of monetary shocks summarised in eq. (55) of section 2. In particular, as originally pointed out in OR, the larger (smaller) the home country – that is, the greater (the smaller) n – the smaller (the greater) the positive impact of a home money increase on its current account. As a result, it is reasonable to expect that smaller and more open countries would show relatively greater effects of monetary shocks on the current account.

Table A.3.3 in Appendix 3 provides two typical indicators of openness to international trade: the ratio of total exports to GDP and the ratio of total trade volumes (exports plus imports) to GDP. Figures refer to average values over the period 1980-1998 for each of the 15 OECD economies. The two ratios allow us to classify the countries in a straightforward manner. In particular, relatively smaller countries seem to be more open to trade (e.g. Austria, Belgium, Portugal and Sweden), whereas low values of the ratios are generally associated with relatively larger economies (for instance, the U.S. and Japan).

To formally test for this prediction, the maximum effect of the estimated current account response to the positive monetary shock in the trivariate SVAR is first calculated. With the only exception of Sweden, Germany and Spain where the peak response is reached between two and four quarters after the shock, as shown in Figure A.3.5, in all the remaining countries the maximum effect is contemporaneous (within the same quarter). It is worth pointing out that, since the variances of the structural disturbances are normalised to one (see section 5.2), all the impulse responses refer to a unit structural shock and, as a result, can be directly compared. Figure A.3.7 in Appendix 3 displays a scatter diagram of the ratio of the overall trade to GDP and the estimated maximum current account response.[43] Although a trend line is not directly included, from the diagram it is possible to see a clear positive correlation between the degree of openness and the sensitivity of the current account fluctuations to nominal shocks.[44]

These results are not only in line with the theoretical prediction of eq. (55) of the *Redux* model, but could also be interpreted as an additional sign demonstrating the

[43] Portugal has not been included because, due to short sample period available, the size of the response, although very high, is very sensitive to the lag length used.

[44] The responses of the real exchange rate and the relative output are much more mixed across countries. Whereas the estimated IRFs are qualitatively consistent with the correct identification of the nominal disturbances, their size does not seem to lead to any clear-cut classification between EMU and non-EMU countries.

consistency of the VAR systems implemented above. At the same time, they are consistent with the idea that current account imbalances are important engines for the recovery of economies that are relatively more open internationally, and provide supportive indications in favour of an asymmetric operativenness of the exchange rate channel of the MTM. Although our estimates are based on the pre-EMU period, in which infra-European nominal exchange rates fluctuations might have played a greater role in the intertemporal trade dynamics, policy-induced exchange rate fluctuations of the Euro versus non-EMU countries could still affect the individual economies in an heterogeneous way, making more open economies relatively more vulnerable to common monetary policy actions of the European Central Bank.

6. Conclusions

Current account fluctuations have always represented not only a cause of concern for policy makers, but also a debated topic amongst academics. Until very recently, most of the literature has focused on productivity and fiscal shocks as the main determinants of the external account. However, since the publication of the Obstfeld and Rogoff (1995, 1996) *Redux* model, there has been increasing theoretical support in favour of a significant role of nominal disturbances. At the same time, many of the subsequent extensions have tended to argue against or at least challenge the original predictions. For instance, some authors have introduced price to market (PTM) behaviour which strongly reduces the potential current account flows between economies resulting from monetary shocks. Others support the latter point on the basis of a low degree of substitution among domestic and foreign goods. These considerations, together with tractability reasons of the underlying theoretical models, have led many authors to completely neglect current account changes in the NOEM.

This chapter aims to provide some new empirical evidence on the role of the current account in the transmission mechanism of monetary impulses by disentangling the fundamental disturbances affecting the external balance with structural VARs estimated separately for fifteen industrialised economies during the pre-EMU period. Before doing this, the empirical relevance of this analysis is motivated through an outline of the monetary version of the sticky-price intertemporal model of Obstfeld and Rogoff (1995, 1996) and a discussion of some extensions introducing PTM and generalising the degree of substitutability between domestic and foreign goods are discussed. These influential models not only offer alternative implications regarding the effects of monetary disturbances on the exchange rate and current account, but also provide a different understanding of the international monetary transmission mechanism. As a result of these competing theoretical setups, the quantitative importance of nominal shocks in generating exchange rate and current account fluctuations is assessed by estimating two SVAR models for each country, in which, similar to Blanchard and Quah (1989) and Clarida and Galì (1994), identification of the structural shocks driving the variables of the system is achieved by imposing previously used long-run neutrality restrictions.

Although the empirical approach implemented in this chapter could be extended in a number of directions to better identify and disentangle additional independent real structural shocks, the results with respect to the effects of nominal impulses are very clear-cut. In particular, the forecast error variance decomposition analysis shows strong evidence in favour of a crucial role of monetary shocks in generating current account movements in most of the industrialised countries under investigation. Moreover, the estimated impulse responses are very much in line with the main predictions of the Obstfeld and Rogoff model, in which unexpected expansive nominal impulses temporarily and significantly improve the current account. These findings clearly reject those class of models featured by a low degree of substitution among domestic and foreign goods. Additionally, it is also found that such effects are quantitatively more important in relatively smaller or more open countries. In contrast with the *Redux* framework in which the PPP always holds and consistent with a number of NOEM models introducing PTM, monetary disturbances generate temporary and significant real depreciations of the exchange rate.

Overall, these findings represent a supportive manifesto in favour of the new class of sticky-price intertemporal models, in which the current account and the exchange rate constitute key variables in the transmission of monetary shocks in open economies. At the same time, although the previous Keynesian theories do not directly focus on the current account determination, the results of this chapter do not reject the operativeness of the traditional exchange rate channel of the international monetary transmission mechanism.

APPENDIX 1 Proofs of the Obstfeld and Rogoff Model

This Appendix provides a deriva tion of some of the main relationships of the Obstfeld and Rogoff (1995, 1996) *Redux* model discusses in section 2.

Proof 1

The demand functions faced by each monopolistic producer are obtained as the solution to the following problem. Maximising eq. (2) of the main text subject to the nominal budget constraint $\int_0^1 p(z)c(z)dz = Z$, where Z is any fixed total nominal expenditure on goods, the Lagrangian function of this problem is:

(A.1) $$L(c(z),\lambda) = \left[\int_0^1 c(z)^{\frac{\theta-1}{\theta}} dz\right]^{\frac{\theta}{\theta-1}} + \lambda\left[Z - \int_0^1 p(z)c(z)dz\right]$$

where λ denote the Lagrangian multiplier associated with budget constraint.

The first-order conditions imply, for all z, $c(z)^{-1/\theta}\left[\int_0^1 c(z)^{\frac{\theta-1}{\theta}} dz\right]^{\frac{1}{\theta-1}} = \lambda p(z)$

For any two goods z and z', therefore, $[c(z)/c(z')]^{-1/\theta} = [p(z)/p(z')]$ or:

(A.2) $$c(z) = \left[\frac{p(z')}{p(z)}\right]^{\theta} c(z')$$

Substituting this expression into the budget constraint leads to:

(A.3) $$\int_0^1 p(z)\left[\frac{p(z')}{p(z)}\right]^{\theta} c(z')dz = c(z')p(z')^{\theta}\left[\int_0^1 p(z)^{1-\theta} dz\right] = Z$$

Using the definition of the price index given in eq. (3) of the main text, both sides of this equation can be divided by P to yield:

(A.4) $$\frac{c(z')p(z')^{\theta}\left[\int_0^1 p(z)^{1-\theta}dz\right]}{\left[\int_0^1 p(z)^{1-\theta}dz\right]^{\frac{1}{1-\theta}}}=\frac{Z}{P}$$

This expression can be simplified to:

(A.5) $$c(z')\left[\frac{p(z')}{P}\right]^{\theta}=\frac{Z}{P} \;\Rightarrow\; c(z')=\left[\frac{p(z')}{P}\right]^{-\theta}C$$

where $C=\frac{Z}{P}$ is the total consumption of the composite good. This last equation implies that demand for good z by home agent j is equal to $c^{j}(z)=\left[\frac{p(z)}{P}\right]^{-\theta}C^{j}$ and, similarly, the foreign agent j's demand is $c^{*j}(z)=\left[\frac{p^{*}(z)}{P^{*}}\right]^{-\theta}C^{*j}$.

Proof 2

Using eq. (8) of the main text to eliminate $p_t(z)$ from the period budget constraint (6) (implying that $p_t(z)y_t(z)=P_t y_t(z)^{(\theta-1)/\theta}(C_t^w)^{1/\theta}$) and dividing the result by P_t, the resulting expression can be plotted into the intertemporal utility function (6). The corresponding unconstraint maximisation problem, therefore, becomes:

$$\underset{y_t(z),\, M_t,\, F_t}{Max}\; U_t=\sum_{s=t}^{\infty}\beta^{s-t}\left\{\log\left[(1+r_{s-1})F_{s-1}+\frac{M_{s-1}}{P_s}+y_s(z)^{(\theta-1)/\theta}(C_s^w)^{1/\theta}-T_s-F_s-\frac{M_s}{P_s}\right]+\right.$$
$$\left.+\frac{\chi}{1-\varepsilon}\left(\frac{M_s}{P_s}\right)^{1-\varepsilon}-\frac{k}{2}y_s(z)^2\right\}$$

where individuals take world consumption C^w as given. The first-order condition with respect to F_t can be written as:

(i) $$\frac{\delta U_t}{\delta F_t}=0 \;\Rightarrow\; -\frac{1}{C_t}+\beta\frac{(1+r_t)}{C_{t+1}}=0$$

which shows how, at a utility maximum, a consumer cannot gain from feasible shifts of consumption between periods. A unit reduction in the first-period consumption, in fact, lowers the utility by $1/C_t$ and the consumption unit saved can be converted (by lending it) into higher second-period consumption that raise utility by $\beta(1+r_t)/C_{t+1}$. The condition (i) states that, at an optimum, these two quantities are equal. From (i) it is possible to obtain the standard consumption Euler equations (for the case where the intertemporal elasticity of substitution σ is 1) expressed by equations (10) and (11) of the main text.

The second first-order condition with respect to M_t is given by:

(ii) $$\frac{\delta U_t}{\delta M_t}=0 \quad \Rightarrow \quad \frac{1}{P_t C_t}=\chi\left(\frac{M_t}{P_t}\right)^{-\varepsilon}\frac{1}{P_t}+\beta\frac{1}{P_{t+1}C_{t+1}}$$

On the left-hand side (LHS) of this condition it is possible to verify that $1/P_t$ is the quantity of current consumption a person must forgo to raise real balances by one unit, and $1/C_t$ the marginal utility of consumption. On the right-hand side (RHS), the first term is the (derived) marginal utility individuals get from having an extra currency unit to conduct transactions. The second term of eq. (ii) is formed by $1/P_{t+1}$, which is the quantity of consumption the individual will be able to buy in period $t+1$ with the extra currency unit, and β/C_{t+1} is the marginal utility of date $t+1$ consumption, discounted to date t. As in the case of the Euler equations, in equilibrium, the terms in the LHS and the RHS have to be equal. From this condition, equations (10) and (11) of the main text and the definition of the home-currency nominal interest rate on date t (i_t) (given by the Fischer parity equation defined in equation (12)), the second FOC becomes:

$$\left(\frac{M_t}{P_t}\right)^{-\varepsilon}=\frac{1}{\chi}\left(\frac{1}{C_t}-\frac{\beta}{(1+i_t)\beta C_t}\right) \quad \Rightarrow \quad \left(\frac{M_t}{P_t}\right)^{-\varepsilon}=\frac{1}{\chi}\left(\frac{i_t}{(1+i_t)C_t}\right)[0,1].$$

The latter equation can be easily rearranged to give the money-market-equilibrium conditions for the home and foreign countries given in eq. (14) and (15) of section 2. The latter conditions equate the marginal rate of substitution of composite consumption for the services of real money balances to the consumption opportunity cost of holding real balances.

The third FOC with respect to $y_t(z)$ is given by:

$$\text{(iii)} \qquad \frac{\delta U_t}{\delta y_t(z)} = 0 \quad \Rightarrow \quad \frac{1}{C_t}\frac{\theta-1}{\theta}\left(C_t^w\right)^{1/\theta} y_t(z)^{\frac{\theta-1}{\theta}} = k y_t(z)$$

which states that the marginal utility of additional revenue earned from producing an extra unit of good z equals the marginal disutility of producing this extra unit of output (implicitly due to forgone leisure). From (iii) the equations (15) and (16) of the main text can be easily derived.

In order to have a full equilibrium, the above FOCs and the budget constraint need also the transversality condition $\lim_{T\to\infty} R_{t,t+T}\left(F_{t+T+1} + \frac{M_{t+T}}{P_{t+T}}\right) = 0$ to be satisfied. The latter is obtained iterating the period budget constraint given in eq. (6) of the main text.

Proof 3

Dividing eq. (6) of the main text by P_t and taking into account the condition characterizing the steady state (e.g. $F_{t-s} = \ldots = F_t = \ldots = F_{t+s} = \overline{F}$ and $M_{t-s} = \ldots = M_t = \ldots = M_{t+s} = \overline{M}$, implying that $\overline{T} = 0$ from eq. (7)) eq. (20) is easily obtained. Equation (21) can be derived from the latter using the intertemporal budget constraint related to the foreign country and (17), according to which $\overline{F}^* = -\left(\frac{n}{1-n}\right)\overline{F}$.

Proof 4

In this special case, the equilibrium is completely symmetric across the two countries implying:

$$\text{(A.6)} \qquad \bar{p}_0(h)/\bar{P}_0 = \bar{p}_0^*(f)/\bar{P}_0^* = 1$$

which holds because, when prices are measured in the same currency, in the globally symmetric equilibrium, any two goods produced anywhere in the world have the same price. Moreover, given this expression, from (20) and (21) it can be easily shown that:

$$\text{(A.7)} \qquad \bar{C}_0 = \bar{C}_0^* = \bar{y}_0 = \bar{y}_0^* = \bar{C}_0^w$$

Substituting these results into the first-order conditions governing each individual's optimal choice of output (15) and (16) leads to:

$$\text{(A.8)} \qquad \bar{y}_0^{\frac{\theta-1}{\theta}} = \frac{\theta-1}{\theta k}(\bar{y}_0)^{1/\theta}\frac{1}{\bar{y}_0} \quad \Rightarrow \quad \bar{y}_0^{1+\frac{1}{\theta}} = \frac{\theta-1}{\theta k}\bar{y}_0^{\frac{1}{\theta}-1}$$

Dividing both sides by $\bar{y}_0^{\frac{1}{\theta}}$, it is possible to obtain:

$$\text{(A.9)} \qquad \bar{y}_0 = \left(\frac{\theta-1}{\theta r}\right)\bar{y}_0^{-1} \quad \Rightarrow \quad \bar{y}_0 = \left(\frac{\theta-1}{\theta r}\right)^{\frac{1}{2}}$$

The same is true for $\bar{y}_0^*$.

Equation (23) can be derived from (13) and (14) by substituting a rearranged version of eq. (19) (which implies that $\bar{r}/(1+\bar{r}) = 1-\beta$) and knowing that inflation is zero in steady state. The latter result also implies that from equation (12) $\bar{i} = \bar{r}$. This can be shown in what follows:

(A.10) $$\frac{\overline{M}_0}{\overline{P}_0}=\chi^{1/\varepsilon}\overline{C}_0^{1/\varepsilon}\left(\frac{\overline{r}}{1+\overline{r}}\right)^{-1/\varepsilon} \Rightarrow \frac{\overline{M}_0}{\overline{P}_0}=\left(\frac{1-\beta}{\chi}\right)^{-1/\varepsilon}\overline{y}_0^{1/\varepsilon}.$$

The same is true for $\overline{M}_0^* / \overline{P}_0^*$.

Proof 5

Eqs. (38) and (39) can be derived by solving for the differences and for population-weighted sums of home and foreign variables. Subtracting eq. (28) from eq. (27), eq. (31) from eq. (31), and eq. (37) from eq. (36), and making use of the PPP equation (24) yield:

(A.11) $$\hat{y}_t-\hat{y}_t^*=\theta[\hat{E}_t+\hat{p}_t^*(f)-\hat{p}_t(h)]$$

(A.12) $$\hat{y}_t-\hat{y}_t^*=-\frac{\theta}{\theta+1}\left(\hat{C}_t-\hat{C}_t^*\right)$$

(A.13) $$\left(\hat{\overline{C}}-\hat{\overline{C}}^*\right)=\left(\frac{1}{1-n}\right)\overline{r}\frac{d\overline{F}}{\overline{C}_0^w}+\hat{\overline{y}}-\hat{\overline{y}}^*-[\hat{\overline{E}}+\hat{\overline{p}}^*(f)-\hat{\overline{p}}(h)]$$

Substituting the steady-state version of eqs. (A.11) and (A.12) into eq. (A.13) gives the following:

(A.14) $$\left(\hat{\overline{C}}-\hat{\overline{C}}^*\right)=\left(\frac{1}{1-n}\right)\overline{r}\frac{d\overline{F}}{\overline{C}_0^w}-\frac{\theta}{\theta+1}\left(\hat{\overline{C}}-\hat{\overline{C}}^*\right)+\left(\frac{1}{\theta+1}\right)\left(\hat{\overline{C}}-\hat{\overline{C}}^*\right)$$

Rearranging the last expression leads to:

(A.15) $$\left(\hat{\overline{C}}-\hat{\overline{C}}^*\right)=\left(\frac{\theta+1}{2\theta}\right)\left(\frac{1}{1-n}\right)\overline{r}\frac{d\overline{F}}{\overline{C}_0^w}$$

From (A.15) it is possible to see that, under the assumption that output were exogenous (implying an inelastic labour supply), a wealth transfer would lead to a steady-state per capita international consumption differential of $\left(\frac{1}{1-n}\right)\overline{r}\frac{d\overline{F}}{\overline{C}_0^w}$. In the

OR model, however, output is endogenous and the impact of wealth transfers on consumption differentials is smaller, because with higher interest income, home agents choose to substitute work for leisure.

It can be shown that combining eq. (A.15) with the relationship $\hat{\bar{y}}_t^w = \hat{\bar{C}}_t^w = 0$, and knowing that for any variable X, $X_t = X_t^w + (1-n)(X_t - X_t^*)$ and $X_t^* = X_t^w - n(X_t - X_t^*)$, eqs. (38) and (39) of section 2 can be easily obtained.

Proof 6

The linearization of the home country's per capita current-account surplus is based on the fact that $F_0 = 0$. This implies that:

$$\text{(A.19)} \qquad F_t = \frac{p_t(h)y_t}{P_t} - C_t$$

By total differentiating this equation and dividing both sides by $\bar{C}_0^w$, it is possible to obtain:

$$\text{(A.20)} \qquad \frac{d\bar{F}}{\bar{C}_0^w} = \frac{\bar{y}_0}{\bar{P}_0\bar{C}_0^w}dp(h) + \frac{\bar{p}_0(h)}{\bar{P}_0\bar{C}_0^w}dy - \frac{\bar{p}_0(h)\bar{y}_0}{\bar{P}_0^2\bar{C}_0^w}dP - \frac{dC}{\bar{C}_0^w}$$

which can be simplified in (46), by making use of eq. (42) and knowing that $dp(h) = 0$ in the short-term. The same is true for (47).

APPENDIX 2   Data Description and Unit Root Analysis

Data

Data have been collected from the International Financial Statistics (IFS) published by the International Monetary Fund (IMF) and the OECD Analytical Database.

REAL EFFECTIVE EXCHANGE RATE (RER): This refers to the CPI-based real effective exchange rate, constructed as a multilateral trade-weighted index According to the IFS definition, a decrease in RER reflects a real depreciation of the corresponding currency.

CURRENT ACCOUNT TO GDP RATIO (CA/Y). The original current account series are in U.S. dollars and non-seasonally adjusted. Following Lee and Chinn (1998, 2002), the national currency-denominated current account is obtained by using the average bilateral exchange rate of each period, and divided by the nominal gross domestic product (GDP). The ratio CA/Y is then seasonally adjusted by regressing it on a series of quarterly dummy variables.

RELATIVE OUTPUT (Y/Y*). This variable is constructed by taking the ratio of the seasonally adjusted home real GDP index (available in the OECD Main Economic Indicators) over the world (OECD) real GDP index. The latter is constructed using PPP exchange rates. An ideal country-specific measure of foreign output could be a trade-weighted average of the real GDP of all the trading partners denominated in the same currency. The problem with such a procedure, however, is that PPP exchange rates are not easily available for most countries. An alternative solution is to convert the national-denominated real GDP into US$ with bilateral exchange rates and calculate a weighted average based on trade weights during the sample period. This approach has been implemented for each country with respect to the G-7 economies, which constitute the main foreign trade partners for most of the OECD-15. Maybe due to nominal exchange rate effects affecting the denominator of the relative output ratio, however, the latter variable appears to perform relatively poorly in the empirical analysis. As a result, the foreign output has been proxied by the OECD GDP index, as in Clarida and Prendergast (1999). This is not by any means a poor measure of world output. If at all, the corresponding ratios dynamics might be biased downwards, above all in relatively larger countries like the U.S. and Japan due to their more substantial role in driving the OECD output index. In order to test for the robustness and the size of the bias, the home real GDP index only has been used as output variable in the trivariate SVAR. While the estimated impulse responses for the real effective exchange rate and the current account remain unaffected, the response of the relative output to the structural shocks slightly increases.

Testing for Unit Roots

Tables A.2.1, A.2.2 and A.2.3 report the Phillip-Perron (PP) and the Augmented Dickey Fuller (ADF) unit root test results for the series RER, CA/Y and Y/Y* respectively. For the real exchange rate series all tests fail to reject the unit root null hypothesis at 1 percent level, whereas the latter is strongly rejected for the first differences. As a result, consistent with a number of empirical papers studying the real exchange rate behaviour in the post-Bretton Woods period, the RER series are assumed to be integrated of order one.

Table A.2.2 shows the results for the current account ratio CA/Y. In most of the countries, the stationarity of our current account series cannot be rejected at 5% significance level. In some cases, namely Finland, Japan and the U.S. (using the PP tests) and France and the Netherlands (with the ADF) the null hypothesis of the presence of a unit root is rejected at 10% only. In general, however, the current account to output ratio can be reasonably assumed to be a stationary variable. These findings are consistent not only with the G7 study of Lee and Chinn (1998, 2002), but also with the recent cross-country paper of Taylor (2002), who cannot reject the stationarity of the CA/Y series in an even wider set of countries.

Finally, Table A.2.3 provides the results for the relative output variable Y/Y*. For all countries, the null hypothesis of non-stationarity cannot be rejected at any level of significance. By implementing the same test for the first differences of the series, the relevant ADF test statistics all reject the non-stationary hypothesis. Therefore, the relative output Y/Y* is assumed to be first-difference stationary.

TABLE A.2.1 Unit Root Tests for RER - Quarterly 1979:2-1998:4

	PP		*ADF*	
	Level	First Difference	Level	First Difference
Austria	-2.56	-7.26***	-2.40	-7.24***
Belgium	-3.02	-5.53***	-2.79	-5.60***
Canada	-1.55	-6.14***	-1.29	-5.93***
Denmark	-3.27	-6.53***	-3.18	-6.61***
Finland	-1.94	-5.75***	-2.12	-5.79***
France	-2.80	-7.25***	-3.17	-7.24***
Germany	-2.87	-6.15***	-2.96	-6.08***
Italy	-1.66	-6.12***	-1.89	-6.09***
Japan	-1.88	-6.16***	-2.20	-6.12***
Netherlands	-3.12	-6.20***	-3.10	-6.23***
Portugal	-1.87	-7.35***	-1.62	-7.37***
Spain	-1.89	-6.23***	-1.64	-6.18***
Sweden	-2.12	-6.88***	-2.33	-6.86***
U.K.	-2.65	-6.70***	-2.77	-6.79***
U.S.	-1.87	-6.22***	-1.54	-6.25***

Notes: PP and ADF denote the Phillip-Perron and the Augmented Dickey-Fuller test statistics for the unit root hypothesis versus the stationarity alternative. Tests for the variable in levels include a constant and a trend. For the first differenced a constant only is included. A truncation lag of three was used in the PP test. The lag length was based on the SW Criterion. (**) and (***) denote significance at 5 and 1% level.

TABLE A.2.1 Unit Root Tests for CA/Y - Quarterly 1979:2 -1998:4

	PP	*ADF*
Austria	-3.35***	-3.85***
Belgium	-5.61***	-5.68***
Canada	-2.89***	-2.96***
Denmark	-3.04***	-2.09**
Finland	-1.86*	-2.06**
France	-4.69***	-3.16*
Germany	-1.94**	-1.86*
Italy	-2.39**	-2.48**
Japan	-1.91*	-2.35**
Netherlands	-4.72***	-2.75
Portugal	-3.97***	-2.00**
Spain	-3.08***	-2.14**
Sweden	-3.57**	-1.83*
U.K.	-2.15**	-1.95**
U.S.	-1.63*	-2.17**

Notes: PP and ADF denote the Phillip-Perron, the Augmented Dickey-Fuller test statistics. For most countries, no deterministic term was included, with the only exception of France, Belgium and the Netherlands where a trend was highly significant. The lag length in the ADF tests was selected on the basis of the SW criterion. (*), (**) and (***) denote significance at 10, 5 and 1% level.

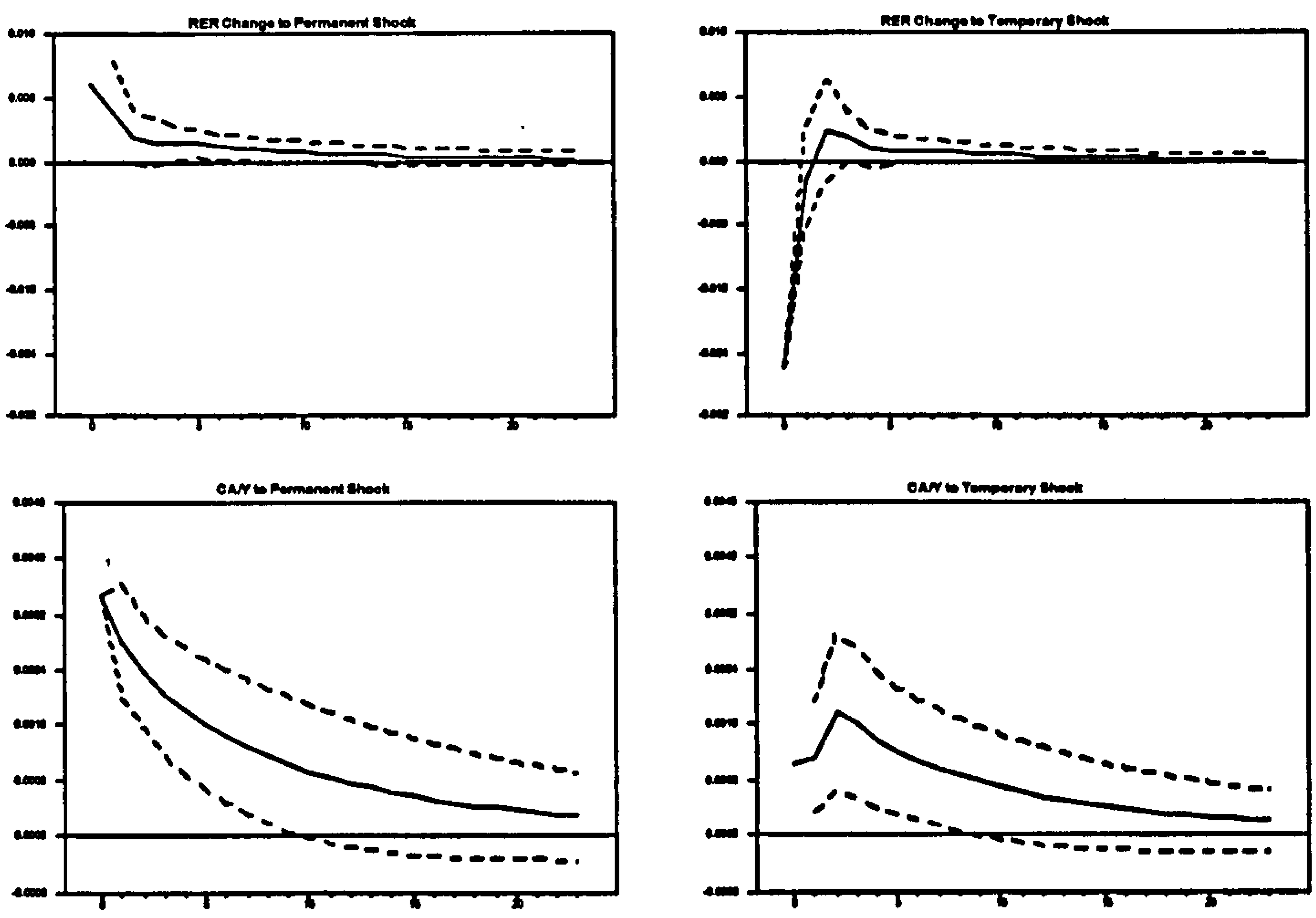
TABLE A.2.3 Unit Root Tests for Y/Y* - Quarterly 1979:2-1998:4

	PP		*ADF*	
	Level	First Difference	Level	First Difference
Austria	-1.97	-7.98***	-1.97	-7.99***
Belgium	-2.05	-5.62***	-2.26	-5.61***
Canada	0.46	-7.32***	0.80	-7.26***
Denmark	-1.53	-5.77***	-1.50	-6.35***
Finland	-1.33	-6.20***	-1.30	-6.14***
France	-2.18	-5.73***	-2.40	-5.74***
Germany	-1.19	-6.61***	-0.88	-6.55***
Italy	-1.22	-5.87***	-1.39	-5.90***
Japan	-1.18	-6.45***	-1.21	-6.50***
Netherlands	-1.48	-6.29***	-1.55	-6.31***
Portugal	-1.86	-4.31***	-2.04	-4.33***
Spain	-1.83	-5.94***	-2.76*	-3.67***
Sweden	-1.20	-6.94***	-1.12	-6.97***
U.K.	-1.58	-6.86***	-1.48	-6.94***
U.S.	-1.19	-6.55***	-1.19	-6.57***

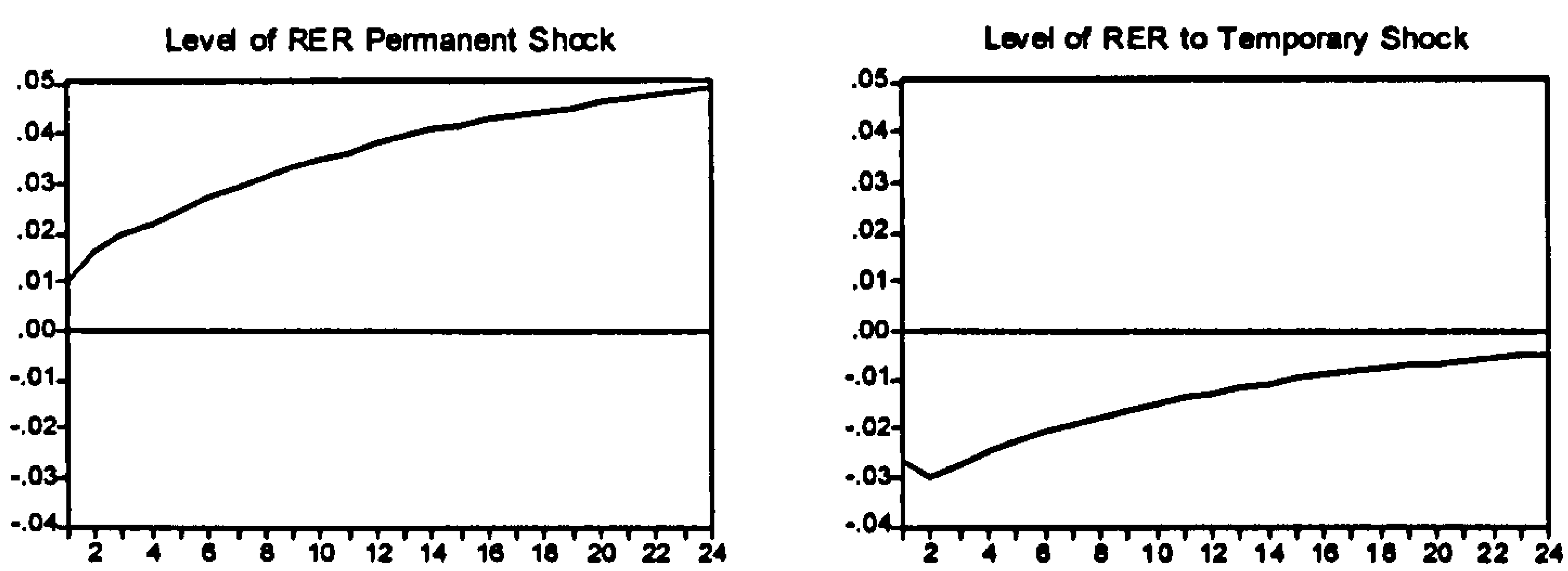
Notes: PP and ADF denote the Phillip-Perron and the Augmented Dickey-Fuller test statistics for the unit root hypothesis versus the stationarity alternative. All tests include a constant. A truncation lag of three was used in the PP test. The lag length was based on the SW Criterion. (*), (**) and (***) denote significance at 10, 5 and 1% level.

FIGURE A.3.1 Impulse Responses for the United States - Bivariate SVAR

Panel A



Panel B



Notes: Panel A shows the impulse responses of the estimated bivariate system in which the real effective exchange rate (RER) is expressed in first differences. Two-standard error confidence bands are constructed using bootstrapping techniques with 1,000 replications. Panel B shows the impulse responses of the level of RER, which are calculated by accumulating the first-difference responses in Panel A.

FIGURE A.3.2 **Impulse Responses of the Ratio of the Current Account to Output - Bivariate SVAR**

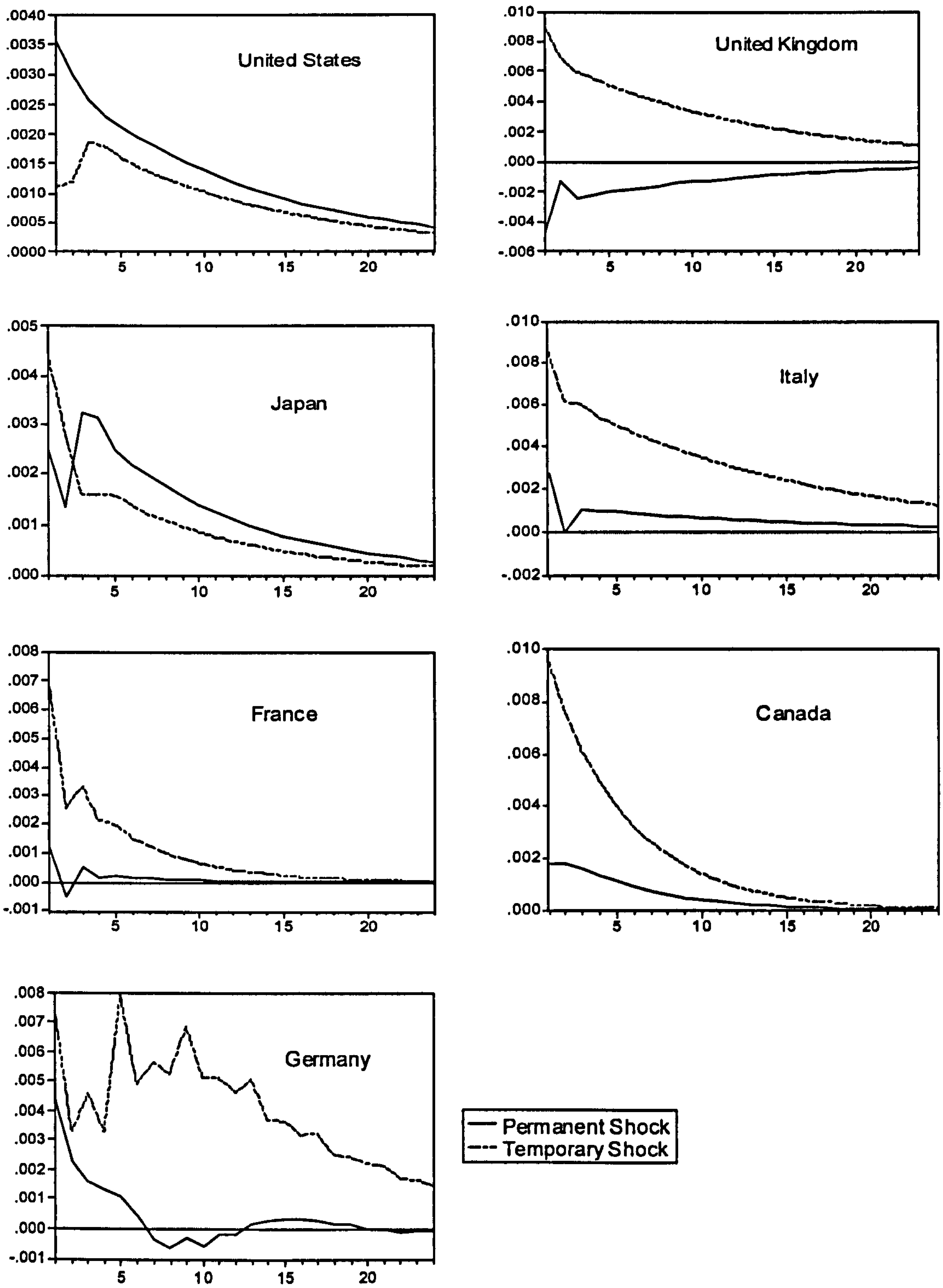


FIGURE A.3.2 Impulse Responses of the Ratio of the Current Account to Output - Bivariate SVAR (continued)

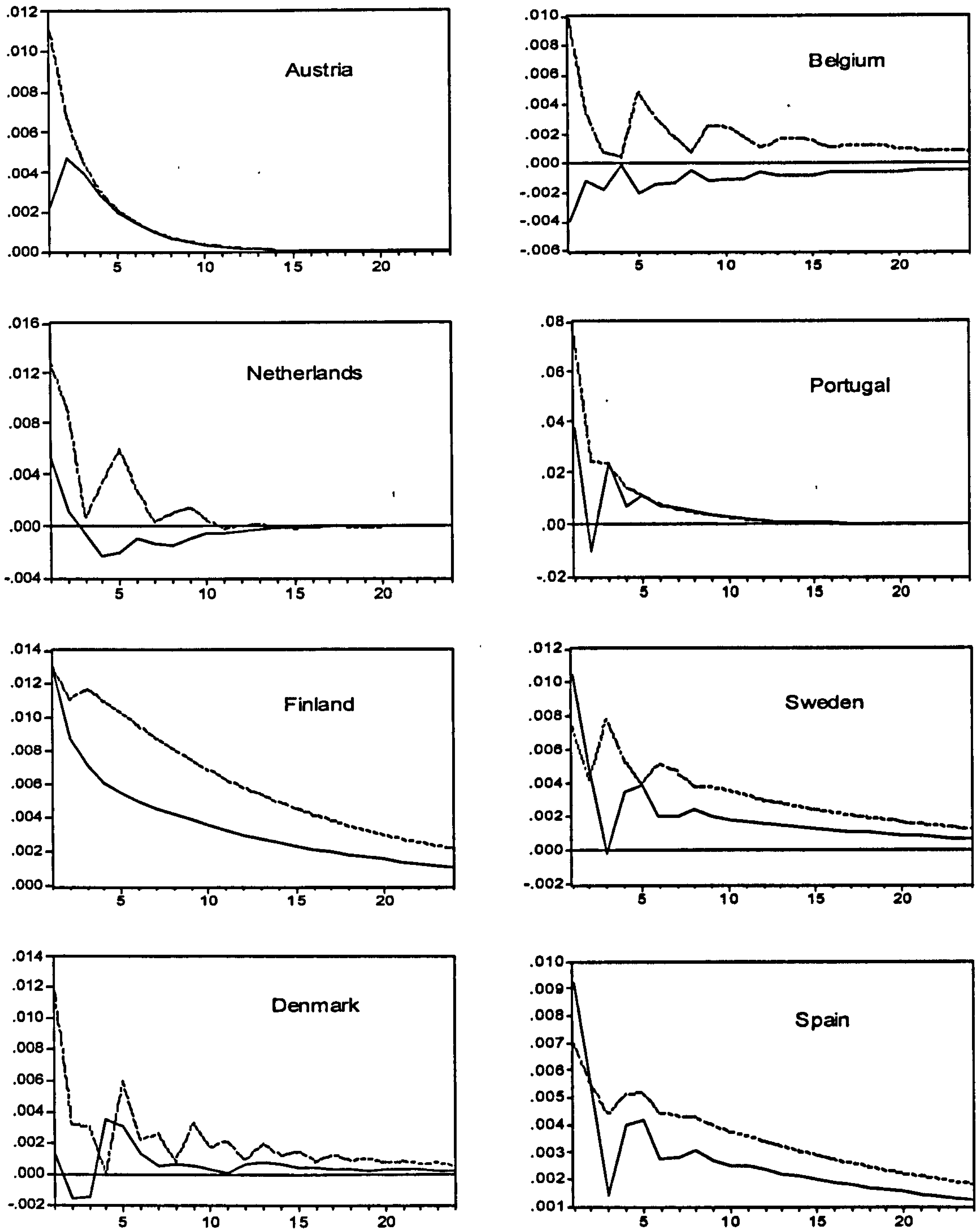


TABLE A.3.1 **Impulse Response Functions of the Bivariate SVAR**

		Permanent Shock	*Temporary Shock*
Austria	RER	**P Appreciation**	**T Depreciation**
	CA/Y	T Surplus	**T Surplus**
Belgium	RER	**P Appreciation**	T Appreciation
	CA/Y	T Deficit	**T Surplus**
Canada	RER	**P Appreciation**	T Depreciation
	CA/Y	T Surplus	**T Surplus**
Denmark	RER	**P Appreciation**	T Depreciation
	CA/Y	T Deficit/Surplus	**T Surplus**
Finland	RER	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	**T Surplus**
France	RER	**P Appreciation**	**T Depreciation**
	CA/Y	T Surplus	**T Surplus**
Germany	RER	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	**T Surplus**
Japan	RER	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	**T Surplus**
Italy	RER	**P Appreciation**	**T Depreciation**
	CA/Y	T Surplus	**T Surplus**
Netherlands	RER	**P Appreciation**	T Depreciation
	CA/Y	T Surplus	**T Surplus**
Portugal	RER	**P Appreciation**	T Depreciation
	CA/Y	T Surplus	**T Surplus**
Spain	RER	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	**T Surplus**
Sweden	RER	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surpl us**	**T Surplus**
United Kingdom	RER	**P Appreciation**	T Appreciation
	CA/Y	T Deficit	**T Surplus**
United States	RER	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	**T Surplus**

Notes: Impulse response functions (IRFs) for the bivariate VARs that include the first difference of (logarithms of) real effective exchange rate (RER) and the level of the ratio of the current account to GDP (CA/Y). IRFs for the former variable were cumulated in order to derive responses in terms of the levels. (T) stands for a tem porary, (P) for permanent change of the *levels* of the relevant variable. See Figure A.3.1 for a representative illustration. Bold entries indicate statistical significance at the 5% level. Standard errors for the impulse responses were constructed using bootstrapping techniques with 1,000 replications.

FIGURE A.3.3 **Forecast Error Variance Decompositions of the Ratio of the Current Account to Output – Bivariate SVAR**

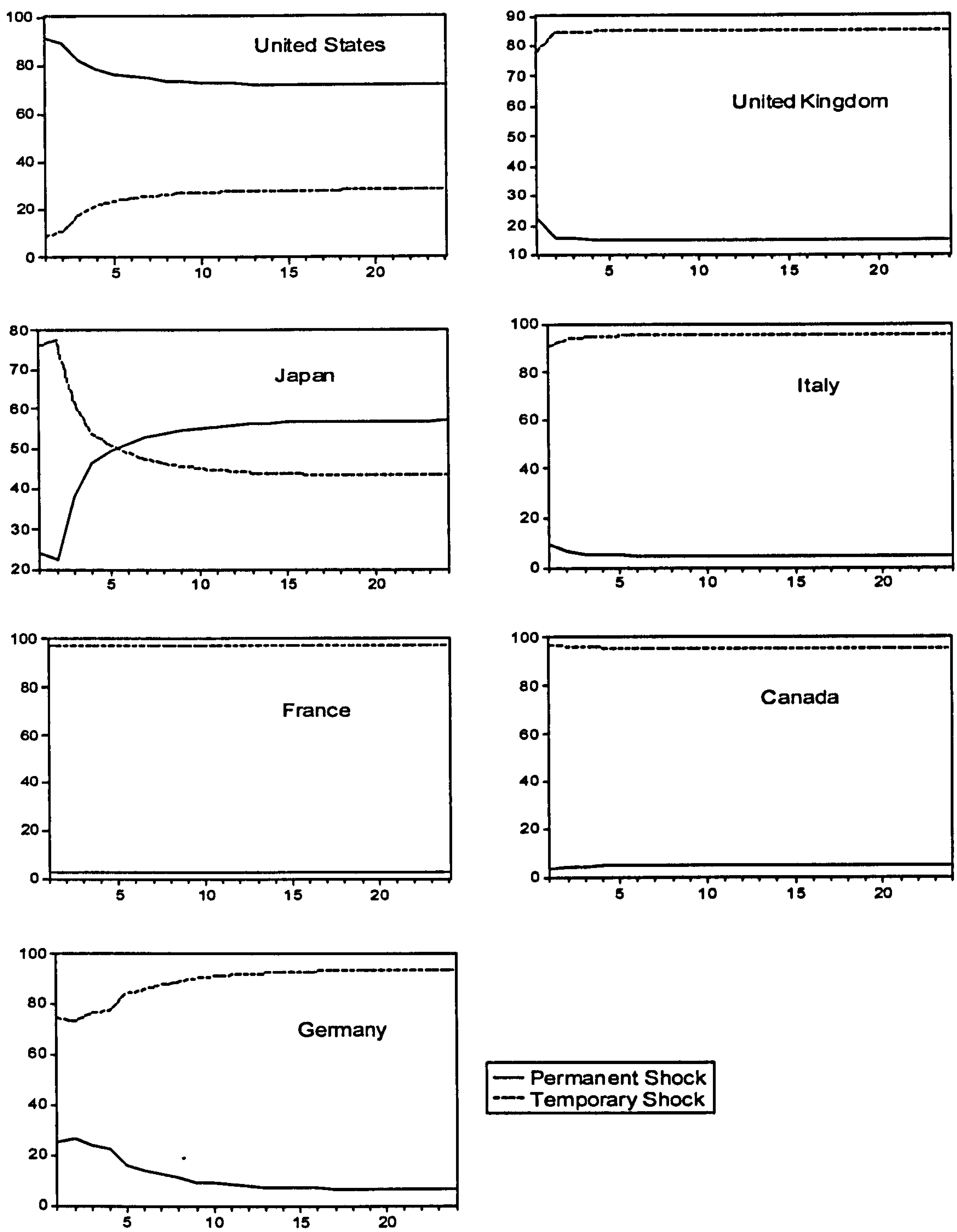


FIGURE A.3.3 **Forecast Error Variance Decompositions of the Ratio of the Current Account to Output - Bivariate SVAR (continued)**

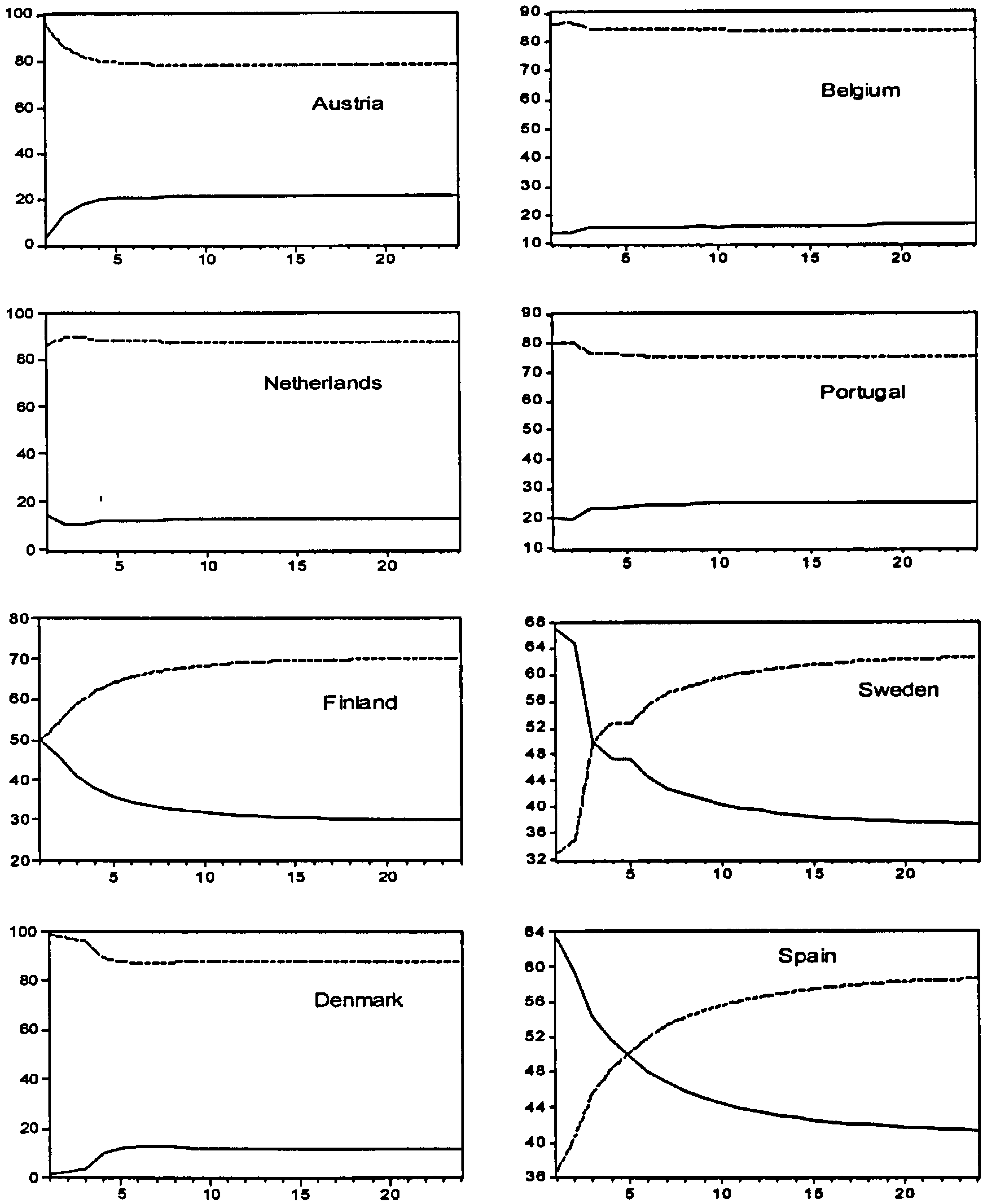
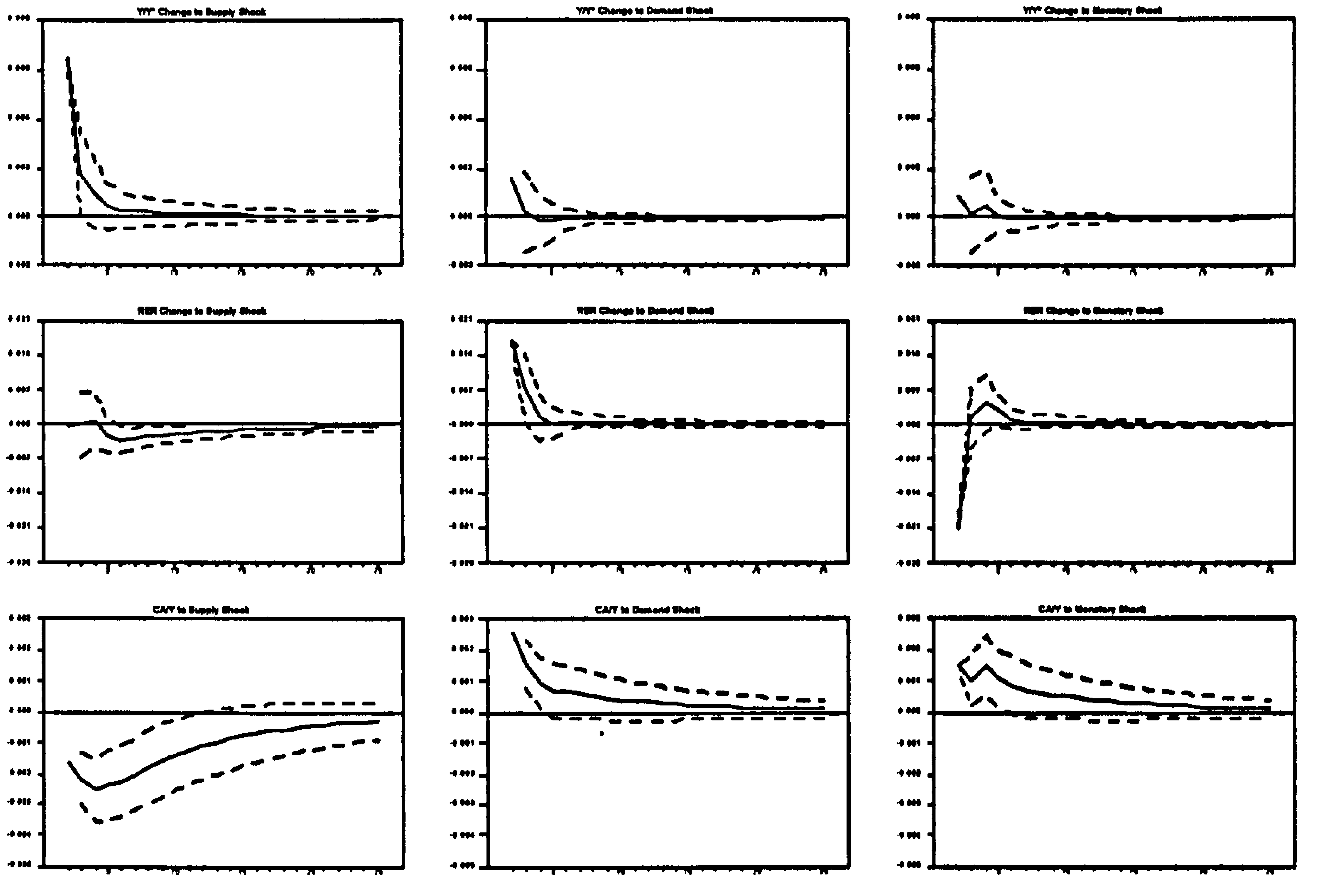
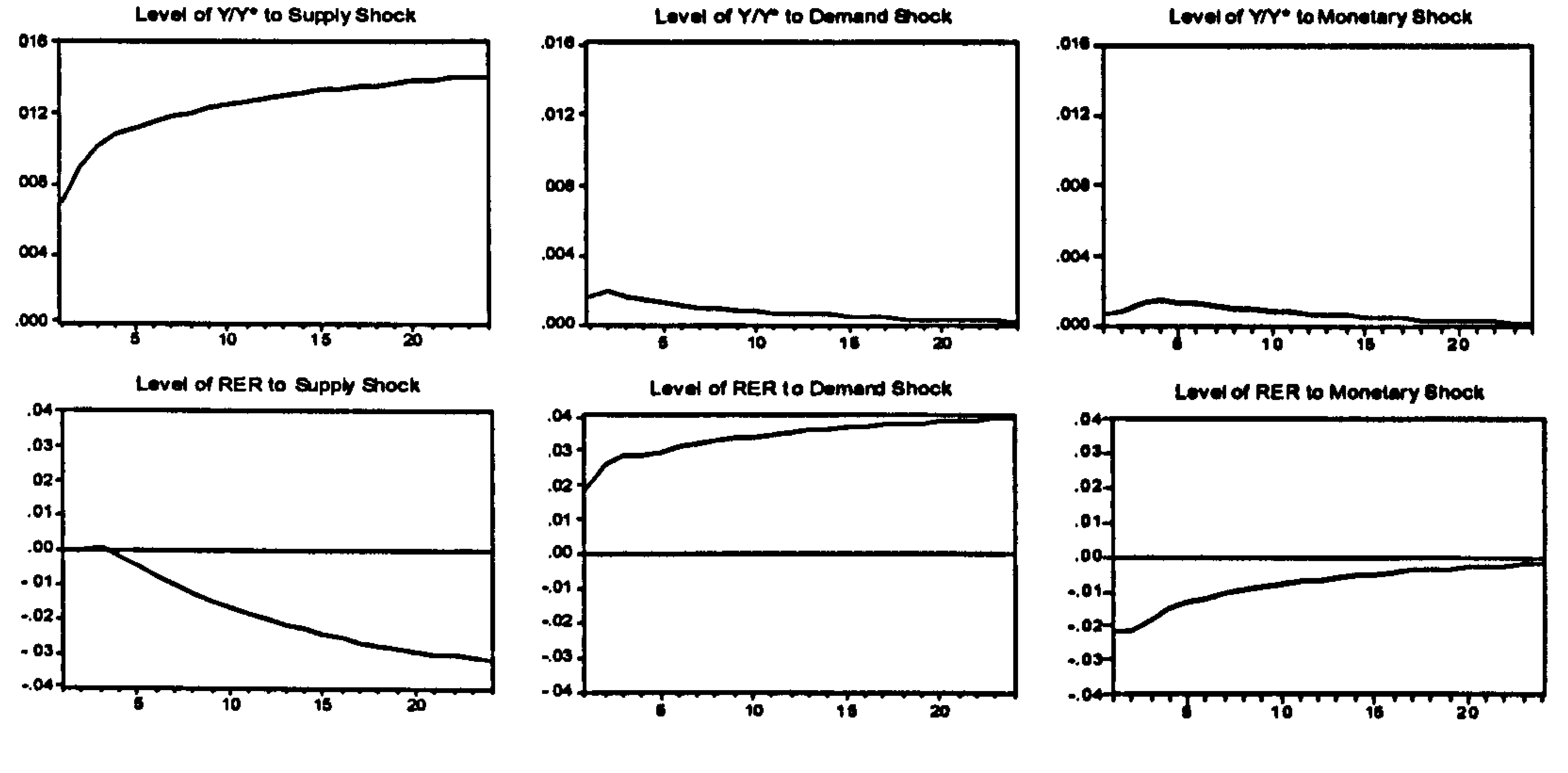


FIGURE A.3.4 Impulse Responses for the United States - Trivariate SVAR

Panel A



Panel B



Notes: Panel A shows the impulse responses of the estimated trivariate system in which relative output and real exchange rate are expressed in first differences. Two-standard error confidence bands are constructed using bootstrapping techniques with 1,000 replications. Panel B shows the impulse responses of the level of Y/Y* and RER , which are calculated by accumulating the first-difference responses in Panel A.

FIGURE A.3.5 Impulse Responses of the Ratio of the Current Account to Output - Trivariate SVAR

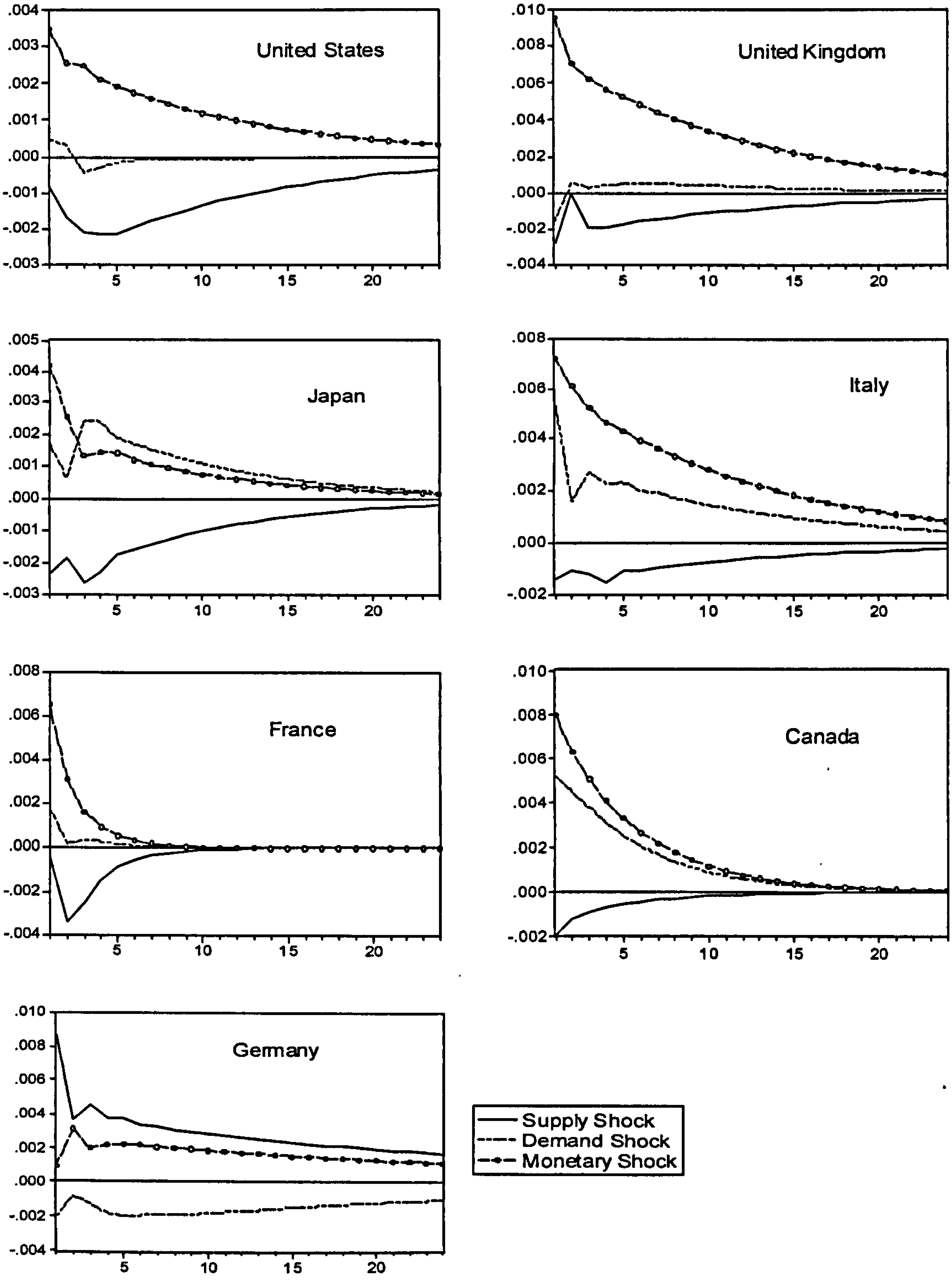


FIGURE A.3.5 Impulse Responses of the Ratio of the Current Account to Output - Trivariate SVAR (continued)

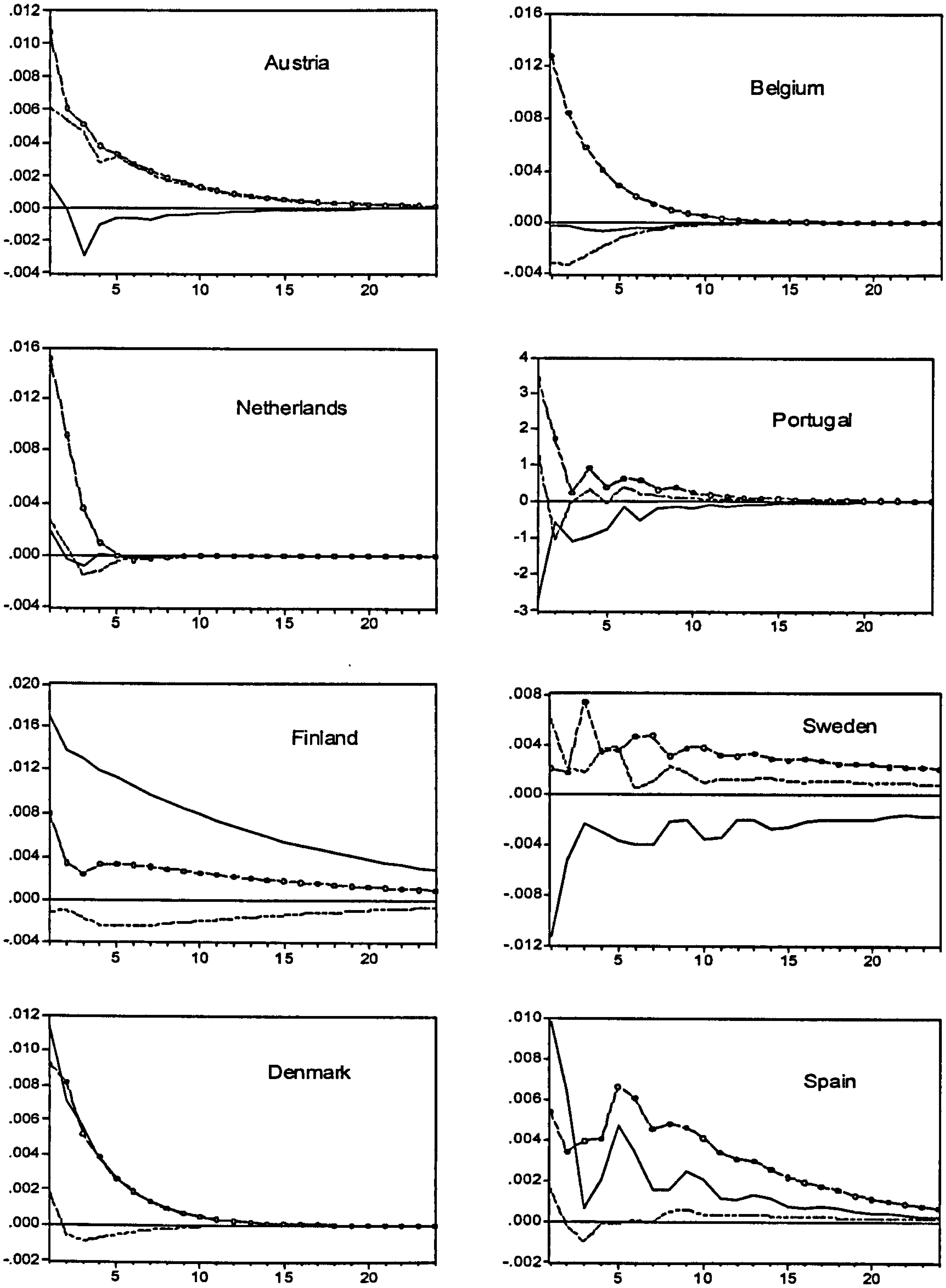


TABLE A.3.2 Impulse Response Functions in the Trivariate SVAR

		Supply Shock	*Demand Shock*	*Monetary Shock*
Austria	Y/Y*	**P Increase**	**T Increase**	T Increase
	RER	P Depreciation	**P Appreciation**	**T Depreciation**
	CA/Y	**T Deficit**	**T Surplus**	**T Surplus**
Belgium	Y/Y*	**P Increase**	T Increase	T Increase
	RER	P Depreciation	**P Appreciation**	T Appreciation
	CA/Y	T Deficit	T Deficit	**T Surplus**
Canada	Y/Y*	**P Increase**	T Decrease	T Increase
	RER	**P Appreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	T Deficit	**T Surplus**	**T Surplus**
Denmark	Y/Y*	**P Increase**	T Increase	T Increase
	RER	P Depreciation	**P Appreciation**	T Depreciation
	CA/Y	**T Surplus**	T Surplus/Deficit	**T Surplus**
Finland	Y/Y*	**P Increase**	T Increase	T Decrease
	RER	P Appreciation	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	T Deficit	**T Surplus**
France	Y/Y*	**P Increase**	T Increase	T Increase
	RER	P Appreciation	**P Appreciation**	**T Depreciation**
	CA/Y	**T Deficit**	T Surplus	**T Surplus**
Germany	Y/Y*	**P Increase**	T Increase	**T Decrease**
	RER	**P Appreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	**T Surplus**	T Deficit	**T Surplus**
Japan	Y/Y*	**P Increase**	T Increase	**T Increase**
	RER	**P Depreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	**T Deficit**	**T Surplus**	**T Surplus**
Italy	Y/Y*	**P Increase**	T Decrease	T Increase
	RER	P Appreciation	**P Appreciation**	**T Depreciation**
	CA/Y	T Deficit	**T Surplus**	**T Surplus**
Netherlands	Y/Y*	**P Increase**	T Increase	T Increase
	RER	P Depreciation	**P Appreciation**	**T Depreciation**
	CA/Y	T Surplus/Deficit	T Surplus/Deficit	**T Surplus**
Portugal	Y/Y*	**P Increase**	T Increase	T Increase
	RER	**P Appreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	**T Deficit**	T Surplus/Deficit	**T Surplus**
Spain	Y/Y*	**P Increase**	T Increase	T Increase
	RER	**P Appreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	**M Surplus**	T Surplus	**T Surplus**
Sweden	Y/Y*	**P Increase**	**T Decrease**	**T Increase**
	RER	**P Depreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	**T Deficit**	T Surplus	**T Surplus**
United Kingdom	Y/Y*	**P Increase**	**T Increase**	**T Increase**
	RER	**P Appreciation**	**P Appreciation**	T Appreciation
	CA/Y	T Deficit	T Deficit/Surplus	**T Surplus**
United States	Y/Y*	**P Increase**	T Increase	T Increase
	RER	**P Depreciation**	**P Appreciation**	**T Depreciation**
	CA/Y	**T Deficit**	**T Surplus**	**T Surplus**

Notes: Impulse response functions (IRFs) for the trivariate VARs that include the first difference of (logarithms of) relative output (Y/Y*) and the real effective exchange rate (RER), and the level of the ratio of the current account to GDP (CA/Y). IRFs for the former two variables were cumulated in order to derive responses in terms of the levels. (T) stands for a temporary, (P) for a permanent change of the levels of the relevant variable. See Figure A.3.4 for a representative example. Bold entries indicate a statistical significance at the 5% level. Standard errors for the impulse responses were constructed using bootstrapping techniques with 1,000 replications.

FIGURE A.3.6 **Forecast Error Variance Decompositions of the Ratio of the Current Account to Output - Trivariate SVAR**

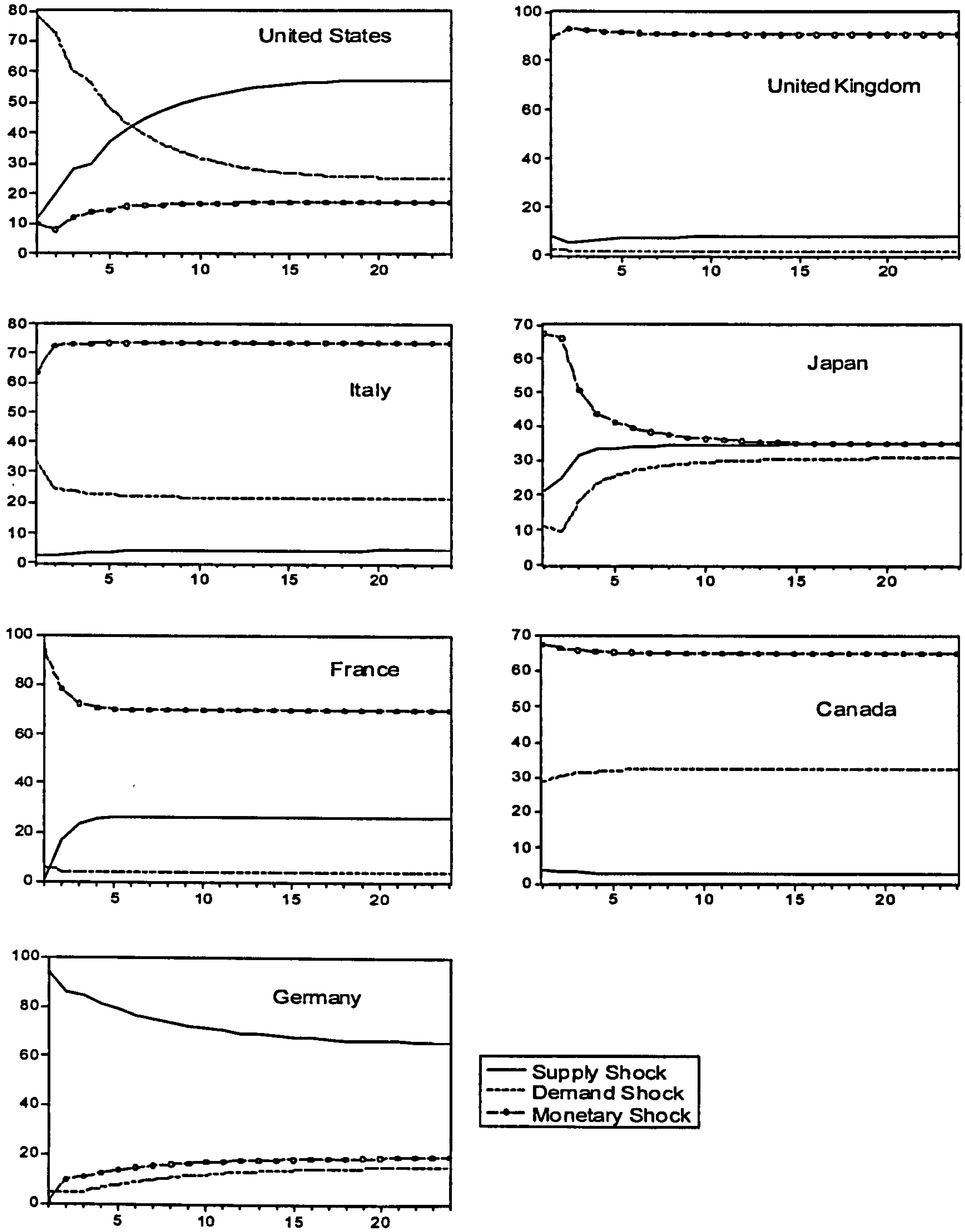


FIGURE A.3.6 Fore cast Error Variance Decompositions of the Ratio of the Current Account to Output - Trivariate SVAR (continued)

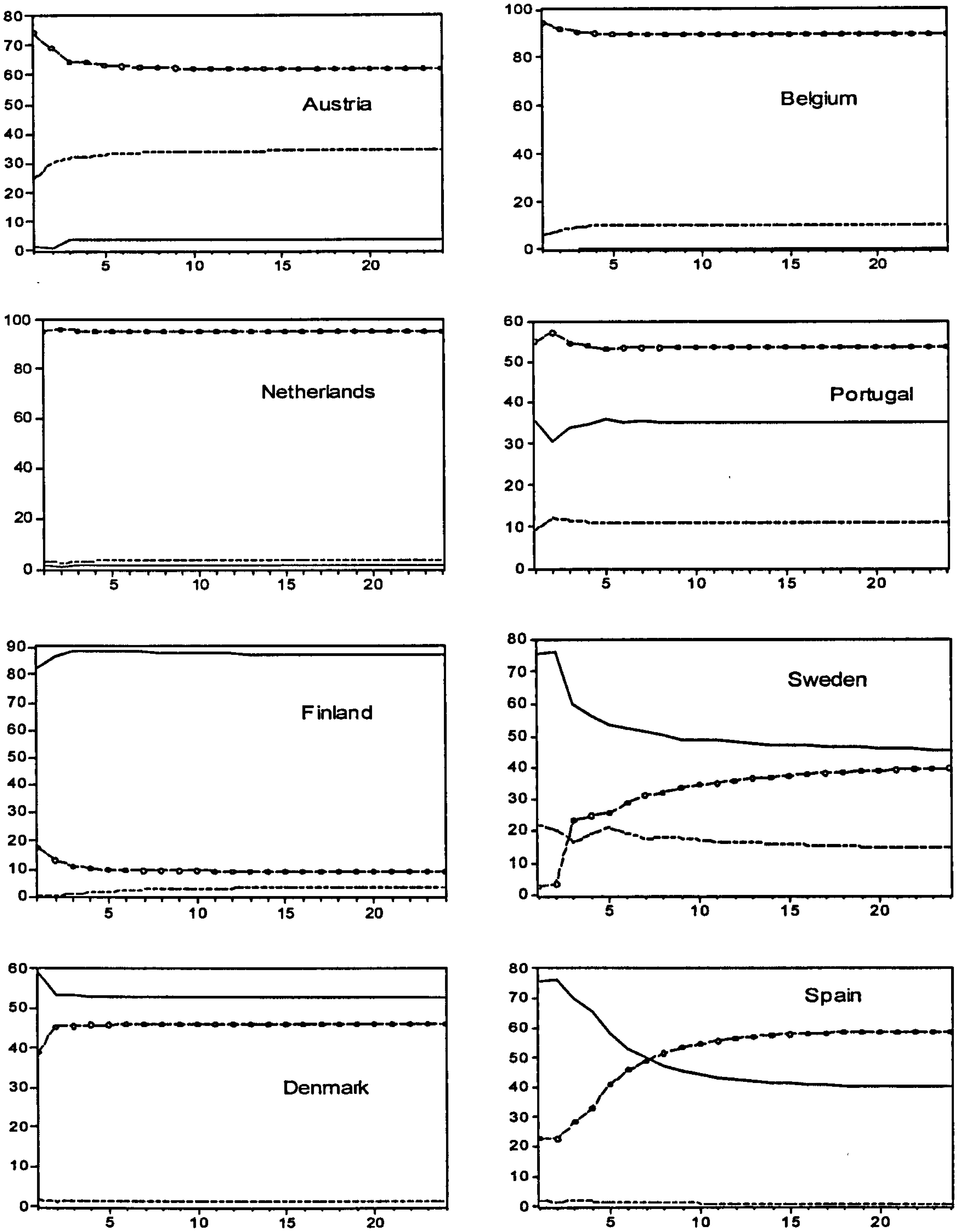
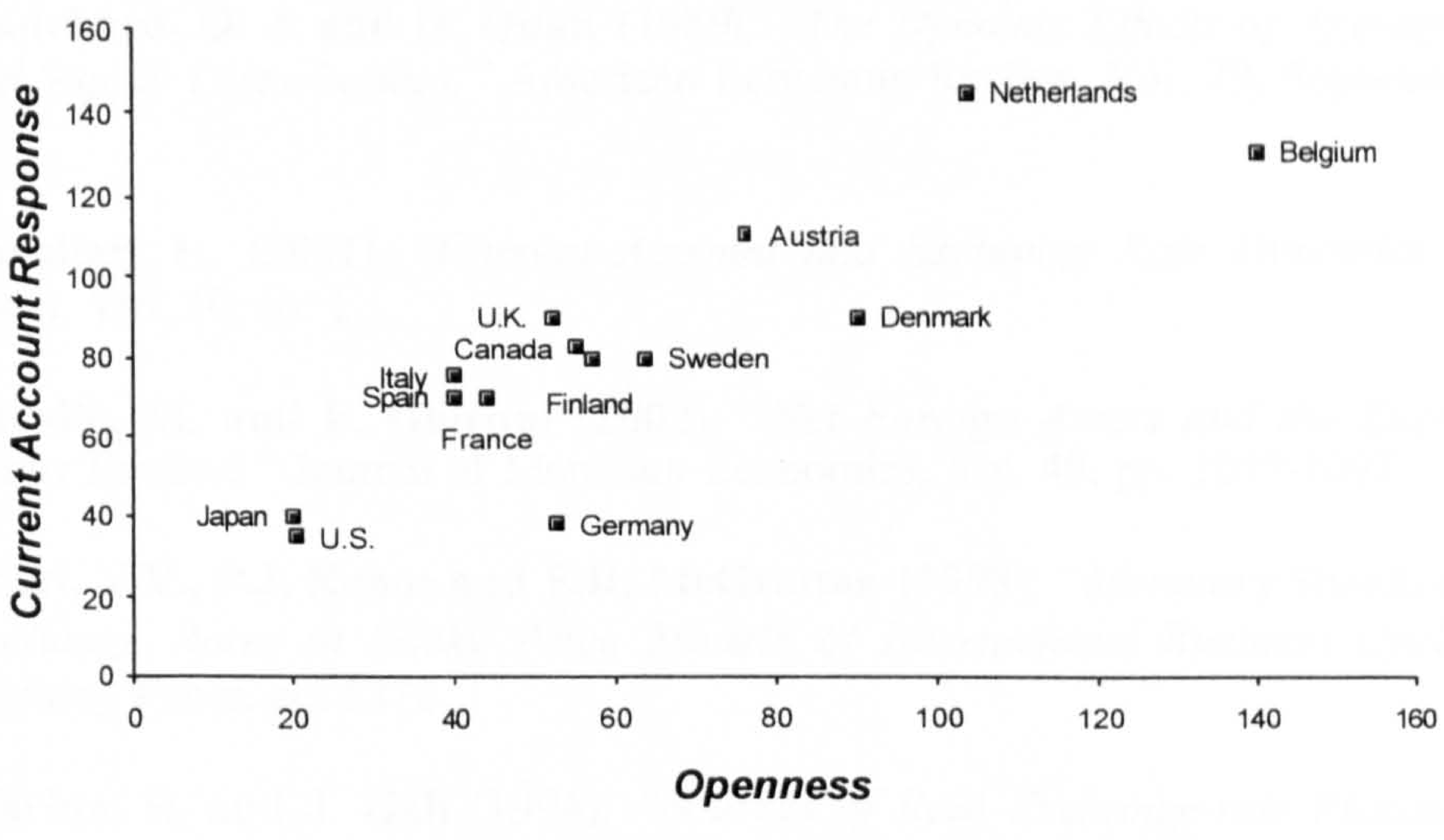


TABLE A.3.3 **Degree of Openness in the 15 OECD**

	Overall Exports of Goods and Service as % GDP	*Overall Exports Plus Imports of Goods and Services as % GDP*
Austria	38.3	76.0
Belgium	71.6	140.3
Canada	27.5	55.1
Denmark	34.9	66.2
Finland	29.1	56.3
France	22.5	44.1
Germany	27.5	52.5
Japan	11.5	20.1
Italy	19.9	39.9
Netherlands	52.2	103.6
Portugal	30.6	69.1
Spain	19.7	40.1
Sweden	33.0	63.5
United Kingdom	25.5	52.0
United States	9.2	20.5
Average	*30.2*	*59.9*
EU Average	*33.7*	*66.7*

Sources: Estimates on OECD data.
Notes: Exports and imports refer to the national income account definition of exports and imports of goods and non-factor services. The figures show average values over the period 1980-1998.

FIGURE A.3.7 **Openness and Current Account Effects to a Unit Nominal Shock**



Notes: The diagram plots the maximum response of the current account to a unit monetary shock versus a trade openness proxy measured as the ratio of totals exports and imports over GDP. Impulse responses refer to the trivariate SVAR. Portugal is not included.

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CHAPTER 2

MONETARY POLICY SHOCKS AND THE ROLE OF HOUSE PRICES ACROSS EUROPEAN COUNTRIES

1. Introduction

The main task of the European Central Bank (ECB) is to conduct monetary policy for the euro area with the primary objective of ensuring price stability and, although not directly specified as a final goal, non-inflationary economic growth. To achieve these objectives, however, the monetary authorities need to have a clear understanding of the effects of their policy upon the economy and the transmission mechanism channels. The size and timing of the monetary transmission mechanism (MTM) can be different across the European Monetary Union (EMU) member countries. Many argue that the monetary regime switch created by the institution of the ECB renders any attempt to study the real effects of national central bank's actions with pre-1999 data less relevant. However, it is doubtful that this institutional change brought about behavioural changes in a sharp and discontinuous fashion (Guiso *et al.*, 2000). In this regard, evidence on agents' reactions to previous regimes can provide some useful insights on how they are behaving, and attempts to throw some light on the MTM based on the past experience might still have informative policy implications.[1]

Over the last few years, many researchers have been trying to compare and analyse the effects of a monetary policy shock across countries.[2] However, while they all try to interpret the results on the basis of the institutional features characterising the individual countries, few of them attempt to identify the relative importance and role of specific channels of the MTM.[3] Even fewer have tried to assess the specific role of asset prices, whose importance in the transmission of monetary impulses have been discussed in Cecchetti, *et al.* (2000) and Mishkin (2001). This is surprising, because the last few decades have witnessed large asset price fluctuations, which have been shown to have affected the real activity in many industrialised countries.[4]

In this field, a particular emphasis has to be given to housing. This is justified by a number of factors. First, housing represents more than half of the net wealth of the private sector in most European economies. Secondly, and differently from financial assets mainly owned by the better-off part of the population, home ownership is distributed more equally across households. Additionally, and perhaps more significantly, housing has a crucial collateral role in the lending sector, which makes it a potential amplifying channel in the real effects of monetary disturbances.

[1] The perfect study would be to analyse the response of various European economies to the same temporal sequence of monetary policy shocks under the new monetary regime of the EMU. However, this is not possible with the newly constituted Monetary Union and the corresponding short data span. As pointed out by Aksoy *et al.* (2002), although the question of how EMU will affect the transmission process is not known, the establishment of the ECB is not a totally new environment, because there has been already a long period of monetary convergence that proceeded it.

[2] See for instance Gerlach and Smets, 1995; Ramaswamy and Sloek, 1997; Barran *et al*, 1998; Dedola and Lippi, 2000; Guiso *et al*, 2000; Ehrmann, 2000 and, more recently, Peersman and Smets, 2001 and Clements, *et al.*, 2001.

[3] An exception in this respect is offered by Clements, *et al.* (2001), in which they try to measure the potency of the credit, the exchange rate and the interest rate channels in the overall MTM.

[4] See Kennedy and Andersen, 1994; Muellbauer and Murphy, 1997; Girouard and Blöndal, 2001; Ludwing and Sloek, 2002 and Case *et al.*, 2002.

In this regard, while some authors have pointed out the special transmission mechanism channels, through which house prices might play a significant role (Maclennan *et al*, 1998; Mishkin, 2001), little empirical evidence has been provided in investigating the quantitative importance of such mechanisms. Moreover, European countries present quite divergent housing and mortgage markets, which reflect differences in the intensity of competition, legal procedures, regulations in the rental sector, fiscal treatment of interest rates, etc. Therefore, under the assumption that these institutional features will converge very slowly (Stephens, 2000), it seems likely that, if asymmetries in the MTM across Euro countries linked to this mechanism are significant, they will last in the medium and long term.

Since the early 1980s, the financial liberalisation process has caused rapid expansion of credit, which determined dramatic house price rises in several industrialised countries. While many authors thought it unlikely that these price developments would recur, the recent general increase of residential prices in most of the European countries shows that, on the contrary, the booms of the 1980s and early 1990s were not a one-off event (ECB, 2003). Given the recent concern that monetary authorities have shown towards asset price dynamics, it seems informative to assess the relative importance of the monetary policy stance in the transmission mechanism operating through real asset values. This might provide not only useful indications in the U.K. and the Scandinavian countries, where the financial liberalisation process has already taken place, but also policy implications for those economies, in which the process is expected to show its effects in the near future.

The aim of this chapter is to provide some qualitative and quantitative evidence on the link between monetary shocks and residential prices and between house prices and economic activity over the EMS period for some European countries.[5] The remainder of the chapter is organised as follows. The next section offers a brief descriptive discussion on the MTM channels, with a particular emphasis on those related to the housing-system. Section 3 presents a concise attempt to identify the main characteristics of the housing and mortgage systems across the EU countries, highlighting the potential implications for house price sensitivity to interest rate changes, and the role of residential asset changes in the transmission mechanism to household non-housing expenditure. Section 4 briefly lays out the econometric methodology, which is based on structural VARs, and provides the empirical results. The first subsection studies the role of monetary shocks to house price fluctuations by making use of impulse response and variance decomposition analysis from alternative SVAR models estimated separately for each of the countries under investigation. The second sub-section explores the role of those (monetary-policy induced) residential price changes on private consumption through the implementation of counterfactual exercises. Section 5 concludes.

2. *The Monetary Transmission Mechanism and the Housing System*

2.1 The Money and Credit Views of the Monetary Transmission Mechanism

In the last few decades, many economists have been focusing on the important question of whether and how a change in the nominal interest rate affects the level of activity in the economy. Whilst a widespread consensus has been reached on the short to medium-run effects of monetary policy shocks on the real activity and long-run effects on the

[5] Due to data availability on high frequency (quarterly) aggregate house price indices, the countries under study are Belgium, Finland, Ireland, Italy, the Netherlands, Spain, Sweden and the United Kingdom.

price level, the debate over the relative importance of monetary transmission channels is still open.

Economists point out a number of different channels, through which a monetary policy change can influence consumption and investment choices. In particular, they distinguish between two main and not mutually exclusive views: the 'money view' and the 'credit view.' The conventional money view refers to the traditional understanding of monetary transmission to real activity based on the standard IS-LM model, in which money supply and interest rate movements directly affect the level of consumption and investment expenditure. Consistent with the Modigliani and Miller (1958) theorem of the irrelevance of the financial tools used for investment decisions, the money view does not make any distinction between bank loans and bond issuing. This aspect is, on the other hand, the key assumption put forward by the credit view which is based on the financial market imperfections approach and on the recognition of the special role played by financial intermediaries in transmitting monetary impulses to output and inflation.

Following the classification provided by De Bondt (2000),[6] the wide spectrum of the monetary policy transmission mechanism can be divided into five different channels, which are shown in Figure 2.1.

The first channel of the traditional monetary view is the direct monetary channel, according to which an increase in the money supply determines a surplus of cash balances, which may cause an expansion of aggregate spending over time.[7]

[6] There are a number of non-technical reports which use similar classifications. See e.g. the ECB (2000, 2002) and the Bank of England (2003).
[7] An additional direct monetary channel operating under flexible prices and wages is the one focusing on misconceptions of the public about aggregate economic conditions, originally developed by Lucas (1972). According to this model, imperfect information on the sources of price changes leads firms to increase

The second, and more important, group of transmission channels is based on the interest rate channels, which can be further classified into:

(i) the cost-of-capital channel;

(ii) the substitution effect channel; and

(iii) the income effect channel.

The cost-of-capital channel emphasises the effect that a rise of the real interest rate, and consequently of the cost of capital, provokes in terms of investment decisions. It is worth stressing that the magnitude of this effect crucially depends upon the ratio between internal and external financing sources, which operators use to finance their investments. Additionally, and more importantly, whereas a change in the official rates unambiguously moves short-term rates in the same direction, the key aspect affecting the operativenness of this mechanism is the reaction of long-term interest rates, which are influenced by the expected path of future interest rates and inflation expectations.[8]

The substitution effect channel points to the role of the opportunity cost of the current consumption expenditure over future consumption (saving). In order for this channel to work fully, agents have to be able to arbitrage completely between present and future expenditure and, at the same time, to choose between different financing sources. In this sense, not only expectations about future real interest rates, but also the presence of liquidity constraints generated by bank practices and the regulatory system affect the possibility of intertemporal trade-off.

The third and last sub-channel is related to the income effect, which captures the role of variations in the stream of net interest payments on the disposable income, and,

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supply by an amount which depends on the relative variability of money supply changes and market specific shocks in the past.

consequently, on the spending decisions. The magnitude and the sign of this channel depend upon several factors, such as the sensitivity of net interest flows to changes in official short-term rates; the maturity structure of the outstanding balance-sheet positions; the distribution between net creditors and net debtors; the distribution between fixed versus variable and short versus long-term interest rate assets and liabilities; the propensity to spend out of disposable income; and, most importantly, the net asset position of firms and households.

As illustrated in Figure 2.1, besides the direct effects of money and interest rates, monetary policy decisions can be transmitted into the real economic activity through asset prices like exchange rates, stock and bond prices, prices of houses and land, etc. In particular, these asset price channels can be split into two main groups: the exchange rate channel on one side, and the Tobin's q and wealth effect channels on the other.

The exchange rate channel of the transmission mechanism emphasises the effects of policy-induced exchange rate movements on net exports and the current account. In the first chapter, this mechanism has been extensively discussed within the new open economy models, and its empirical operativeness has been shown for a number of industrialised countries with different degrees of trade openness.

As far as the other asset price channels are concerned, the Tobin's *q* effect[9] refers to a situation in which a high value of q implies that companies can issue stock at favourable conditions compared to the costs of new plants and investments. In this sense, investment is attractive. If, on the other hand, the value of *q* is low, new investment spending will be discouraged.

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[8] The role of inflation expectations and credibility are discussed below.

[9] As defined by Tobin (1969), "*q*" is the market value of the firms divided by the replacement costs of capital.

Similarly, the wealth effect channel refers to the effects that asset price fluctuations produce on the wealth of the economic agent holding those specific assets. It is clear that a crucial aspect affecting the potency of this effect is related to the liquidity or 'perceived' spendability of these assets. While financial assets are relatively more liquid that real assets, the deregulation process, started in the 1980s, has brought dramatic changes in the housing and mortgage markets, which have made housing much more spendable. This point will be further developed in the following sections.

So far the traditional monetary view of the transmission mechanism has been discussed. However, more recently additional channels based on the credit view have been explored.[10] As emphasised by Bernanke and Gertler (1995), this view focuses on financial market imperfections and, in particular, on frictions in the credit markets, which affect the degree of substitution between different sources of financing. Bernanke and Blinder (1988) indicate two main effects through which credit enter the MTM: the bank lending channel and the balance sheet channel.

The bank lending channel emphasises the special role that banks play in alleviating the problems of asymmetric or incomplete information in the credit market and in providing finance sources to certain borrowers (typically households and small firms), which depend heavily on bank loans. As a result of that, a monetary tightening, which drains deposits from the banking system, puts banks in need of adjusting their portfolios by reducing their supply of loans. Banks then increase their lending rate or reduce their loans by rationing the least creditworthy loan applicants. These two possible outcomes decrease the agents' economic activity and the consumption and investment expenditure.

[10] See the seminal paper of Bernanke and Blinder (1988) and Stein (1995).

The balance sheet channel focuses on the potential effect of monetary policy on borrowers' financial position or net worth linked to their income accounts and balance sheets. According to this channel, a monetary shock, in affecting the borrowers' net worth, also modifies the overall terms of credit and, consequently, the external finance premium (e.g. the spread between the costs of self-financing and external credit). In particular, an expansionary monetary policy (i.e. a reduction of interest rates) is expected to improve the cash flows of firms, and lead to a rise in house, equity and other asset prices. Both factors strengthen borrowers' net worth and, given that most of the credit to the private sector (and in particular households) is provided on the basis of asset-backed collateral, this might feed further increases of loans. The balance sheet channel plays a crucial role in the MTM and many authors agree that the financial liberalisation process of the 1980s and 1990s might have strengthened its relative importance.

It should be made clear that the credit view is not an alternative to the money view, but an additional way through which monetary policy actions affect private spending and investment. In fact, as argued in Dale and Haldane (1993), it could either increase the potential effects of monetary policy if banks lending rates move more than one-to-one with changes in the official money market rates, or decrease it if they respond sluggishly to movements in the market rates.[11] Clearly, the significance of this channel depends upon several conditions such as the competitiveness of the banking system, the importance of the bank credit to finance investment and consumption, the health of the banking system, etc.

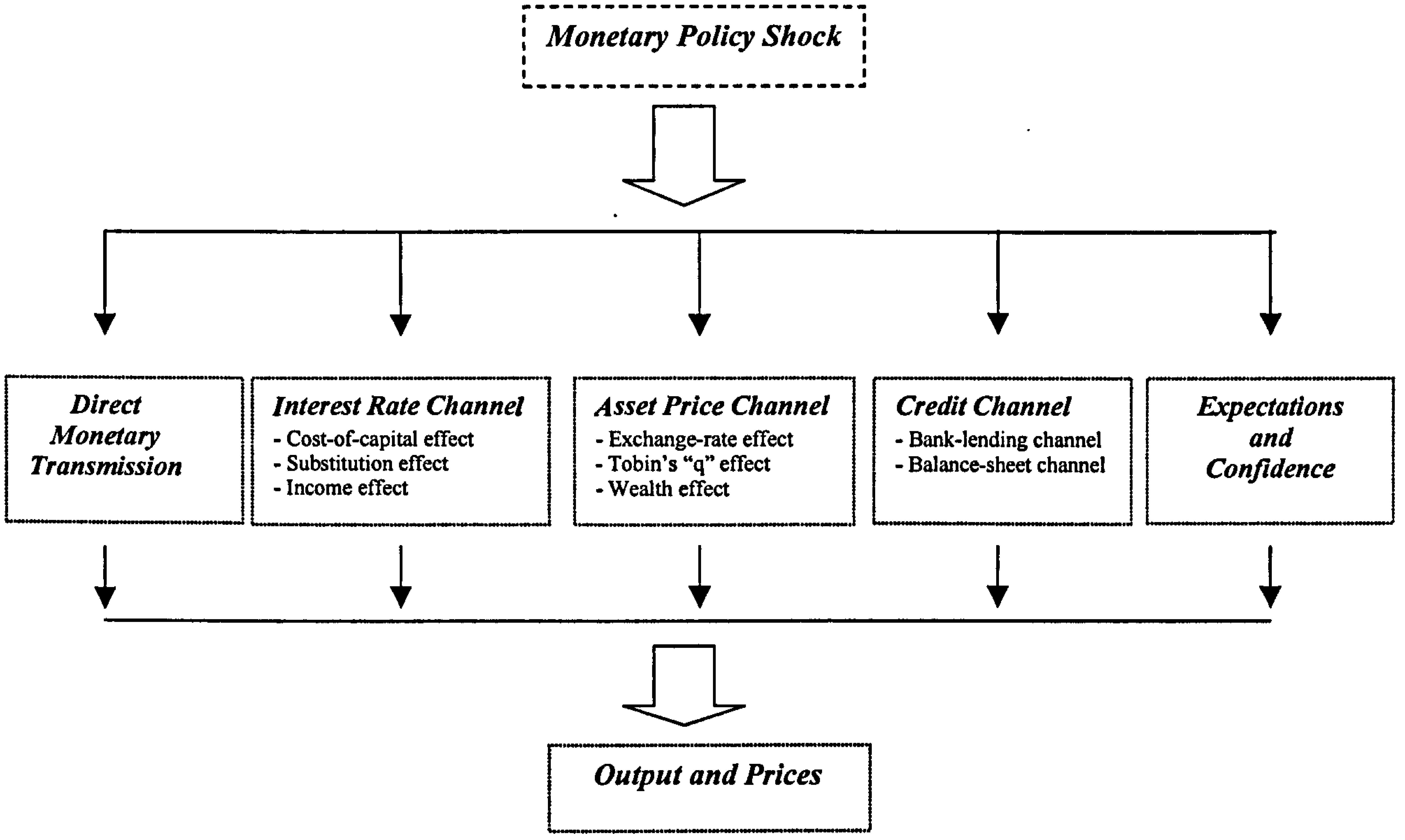
[11] Several authors have researched on the lending interest rates responses to changes in interest rates controlled by the monetary authorities. The pass-through from market rates to bank interest rates is discussed below.

Finally, Figure 2.1 shows the last monetary transmission channel, which operates through expectations, confidence and uncertainty. In this respect, monetary actions can affect expectations about the future course of the real economy and the confidence with which they are held (Bank of England, 2003). In particular, according to the monetary authorities' credibility and ability to influence financial markets and the whole economy, monetary policies can be more or less effective through changes in expected labour income and wages, unemployment, profits, etc.

The direction of these expectation effects, however, is difficult to anticipate, due to a large degree of uncertainty. For instance, agents might interpret a restrictive monetary policy as a sign that the economy is growing faster than expected, boosting their confidence. At the same time, an anti-inflationary policy could also lead to lower expectations of future growth. Uncertainty also makes it harder to distinguish between good and bad credit risks and makes adverse selection and moral hazard problems more severe, causing difficulties in raising financial sources for investment and consumption. The latter aspects vary from time to time, according to the specific policy regime and the general level of consumer's confidence.

It is important to stress the fact that the basic classification displayed in Figure 2.1 does not imply that these channels are not closely interrelated. On the contrary, not only are there important interactions between variables, but the figure does not include potential 'feedback mechanisms' between real economy and monetary policy stance, which, in reality, are central to the conduct of monetary policy.

FIGURE 2.1 The Monetary Policy Transmission Channels



Sources: Elaboration from De Bondt (2000).
Notes: For simplicity, this figure does not display all interactions between variables.

2.2 *The Monetary Transmission Channels Related to the Housing System*

Although the money and credit views provide different explanations of the functioning of the MTM, they share a common understanding of the importance of asset prices and, in particular, house prices. This is justified by a number of factors, such as the quantitative importance of housing wealth in most of the industrial economies, as well as the special characteristics of the housing system in the lending sector. Residential assets, however, are very different from other tangible and financial goods, in the sense that they have the dual nature of commodities yielding utility and investment assets (Miles, 1995). Additionally, housing markets are characterised by a number of special features, which imply a complex role in the transmission of macroeconomic shocks.[12] It is probably the presence of such factors, which makes it difficult to compare housing systems across countries and, more importantly, provide a clear-cut classification of the channels through which these markets enter the MTM. What follows tries to provide a simple attempt, which can help to identify specific 'housing-system-related' channels.

Maclennan (1994) argues that an interest rate decrease has three main first-round effects on the housing sector, which might imply a direct or indirect role of the housing market in the transmission mechanism of monetary policy:

(1) increase of construction of new dwellings and renovation of existing dwellings;
(2) existing borrowers with variable rate mortgages will have a smaller share of their disposable incomes being absorbed by mortgage interest costs;
(3) first-time buyers and likely-moving owners may be encouraged to purchase.

[12] Miles (1995) indicates some special characteristics. They can be summarised as follows: durability, uniqueness and heterogeneity, short-run rigidity of supply, the possibility of raising loans against housing collateral for finance house and non-house purchases, the existence of a well developed secondary market, significant price volatility, housing wealth as the most important component of the value of estates bequeathed, tax treatment, large involvement of financial intermediaries in the housing market and the presence of a closely related rental housing market.

(1) According to the traditional cost-of-capital channel, a decrease of the real interest rate affects investment decisions on the basis of whether the cost of borrowing will be more than covered by the rate of return on capital. Whereas some authors argue that interest rates have their effects on investment above all (and possibly 'only') through expectations of future returns and profits, several empirical papers provide strong evidence in favour of considerable and powerful influences running from short-term interest rates or cost of borrowing to residential investment.[13]

(2) The second effect is given by the interest-rate-income effects, according to which the existing borrowers with variable mortgage interest rates will take advantage of the lower cost of borrowing (and higher disposable income) by increasing their consumption and investments. Crucial factors in this context are the mix between fixed and variable mortgage interest rates and the size of the outstanding mortgage debt.

(3) Finally, lower interest rates are likely to stimulate housing demand of first-time buyers and likely-moving owners. Given the rigidity of real estate supply in the short run (due to restricted land zoning and planning policies), they will also impact on house prices. Obviously, the sensitivity of the demand for housing to a short-term interest rate crucially depends on the pass-through from official to bank mortgage interest rates. To this regard, different institutional characteristics of the banking system might play a significant role. Additionally, the mix of flexible versus fixed interest rates on mortgages might be significant. In particular, for countries with a high diffusion of long-term fixed rates, monetary policy shocks will be propagated much more slowly and, therefore, housing demand will react with some delay.[14] Additionally, similar to

[13] See Bank of England (1990), Bernanke and Gertler (1995), Oxley and Smith (1996), Barran *et al.* (1998) and MacCarhy and Peach (2002) for some empirical evidence.

[14] As emphasised in the previous section, the reaction of long-term rates to policy-controlled interest rates are affected by inflation expectations and future short-term rates. While the mix between fixed and

other asset prices (e.g. bonds and equities), house price increases reflect a rise in the present value of future streams of housing services. Other effects may come from increased income expectations, expectations of future house price rises as well as speculative behaviour (ECB, 2003).

Changes in residential prices followed by shifts in housing demand are thought to have significant effects on economic activity through a number of channels. First, house price fluctuations can affect residential constructions. In particular, on the basis of the so-called Tobin's q theory, when the ratio between house prices and construction costs is above unity, it is profitable for agents (i.e. individuals and construction companies developers) to build new dwellings. The responsiveness of supply of new housing to house price movements (as well as interest rates and other demand shocks), depends upon the degree of competition in the construction industry, building regulations, land planning, availability of specialised labour and fiscal treatment of new housing. The ECB (2003) and Ball (2002) review some of the main cross-country studies, from which it emerges that a relatively low price elasticity of supply exists in Europe by comparison to the U.S., above all in the short run.

Second, residential price movements might lead to important income implications deriving from the market rented sector. Higher house prices cause higher rents for tenants. Higher revenues for landlords or institutional investors owning rental housing can partially offset the negative income effects that tenants face. Under the assumption of a higher marginal propensity to spend out of income for the latter agents, however, it is reasonable to expect an overall negative income effect. The strength of

variable rates reflect differences in tastes and traditions, it is also true that individual inflation histories have played a crucial role. The new environment characterised by low and homogeneous inflation rates across countries will favour a progressive convergence towards fixed rate instruments.

this channel depends upon the housing tenure structure, the functioning of the rental market and the different reactions of agents (tenants, landlords and institutional investors).

Third, a rise in house prices is likely to have positive saving effects for households planning to purchase a house, above all in those countries with high down-payment requirements or with a poorly developed (and rationed) housing finance system, which implies a higher use of internal funds (Kennedy and Andersen, 1994 and Muellbauer and Lattimore, 1995). The strength of this effect depends upon the required deposit/value ratios, which can be quite low in deregulated financial systems and highly competitive lending markets. Positive saving (or substitution) effects are also operative through changes in imputed rents of home-owners, although their perception of these higher housing costs may be relatively low.

The other key and potentially more effective mechanism through which house price changes can affect the real activity is related to the balance-sheet channel, according to which home-owners are able to borrow against the (rising) collateral values. Miles (1994) and Muellbauer and Murphy (1997) have emphasised the role of housing equity withdrawal (that is, the excess of net - of repayment - new lending for house purchase over all forms of investment in residential property), which, in the financial liberalisation process of the 1980s, provided households with liquidity to use in non-housing consumption in the U.K.. The ability of house-owners to extract equity embodied in housing wealth depends upon the competitive conditions in the mortgage markets, which affect the average loan-to-value ratios, ease in re-mortgaging, possibility of second mortgages, and, in general, a greater availability of mortgage products.

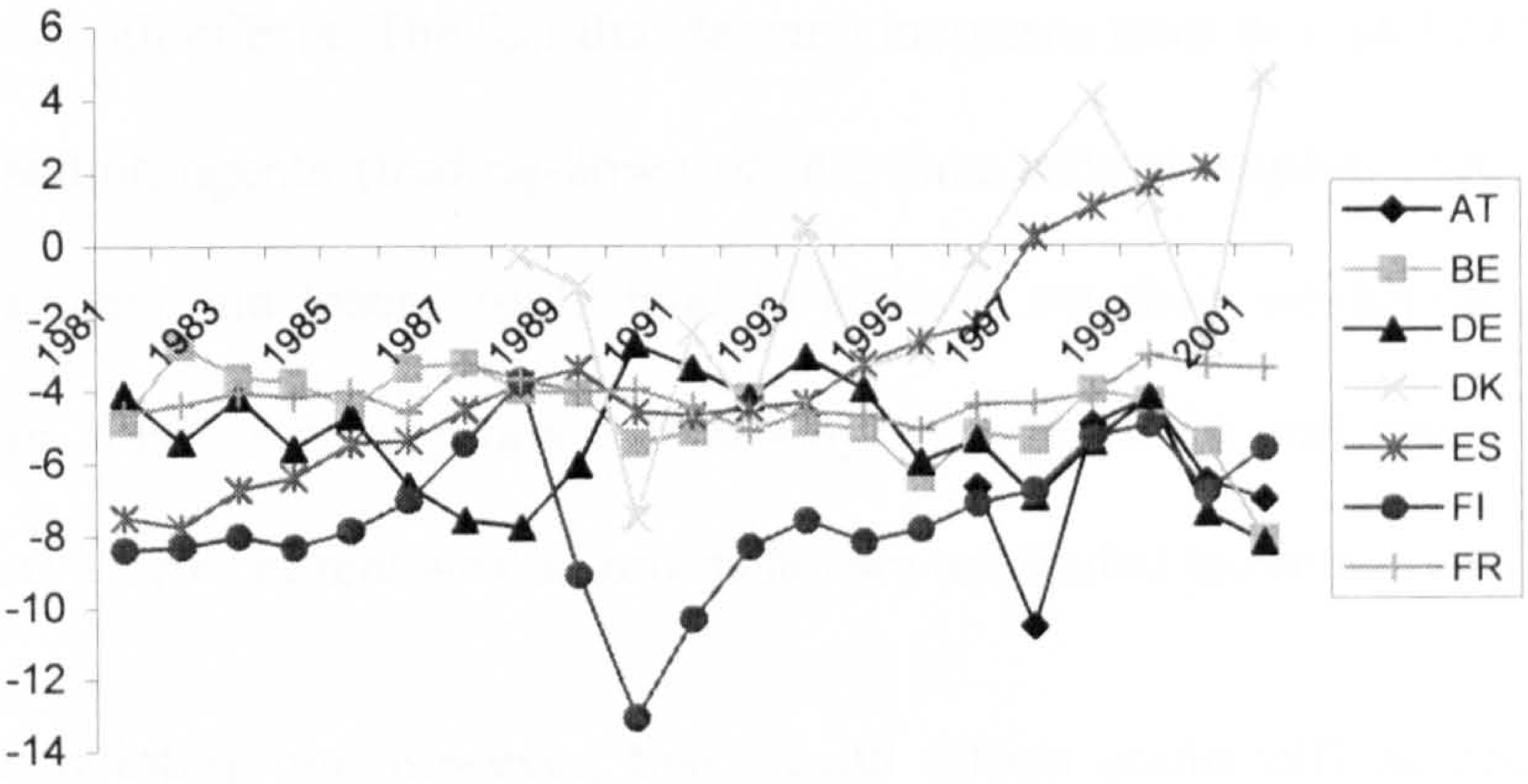
Figure 2.2 shows an estimate of the mortgage equity withdrawal (MEW) in the European economies during the 1980s and 1990s.[15] This measure is defined as the gross change in nominal mortgage debt minus residential investment as percentage of household disposable income. The deregulation of the mortgage market in the 1980s led to a significant equity withdrawal in the U.K. and Sweden. In both countries this was followed by a period of housing equity injections which were reverted during the late 1990s. In the remaining European economies there seems to have been a tendency for housing borrowing to exceed the residential investment for most of the sample period. The only exceptions are Denmark, Spain and the Netherlands in which, during the second half of the 1990s, some housing equity withdrawal occurred. These figures are consistent with high mortgage debt levels present in most of the above countries and a wide diffusion of housing finance instruments, which, following the regulatory changes of the last two decades made it easier for households to withdraw house equity.[16] Additional characteristics of the mortgage markets which partly explain the use of MEW are further analysed in the next section.

[15] The ECB (2003) (Table 5.1, last column), in summarising some of the latest features of the EU mortgage systems, shows that some countries (Denmark, Germany, Ireland, the Netherlands, Finland, Sweden and the United Kingdom) offers instruments that permit households to tap on their housing wealth directly.

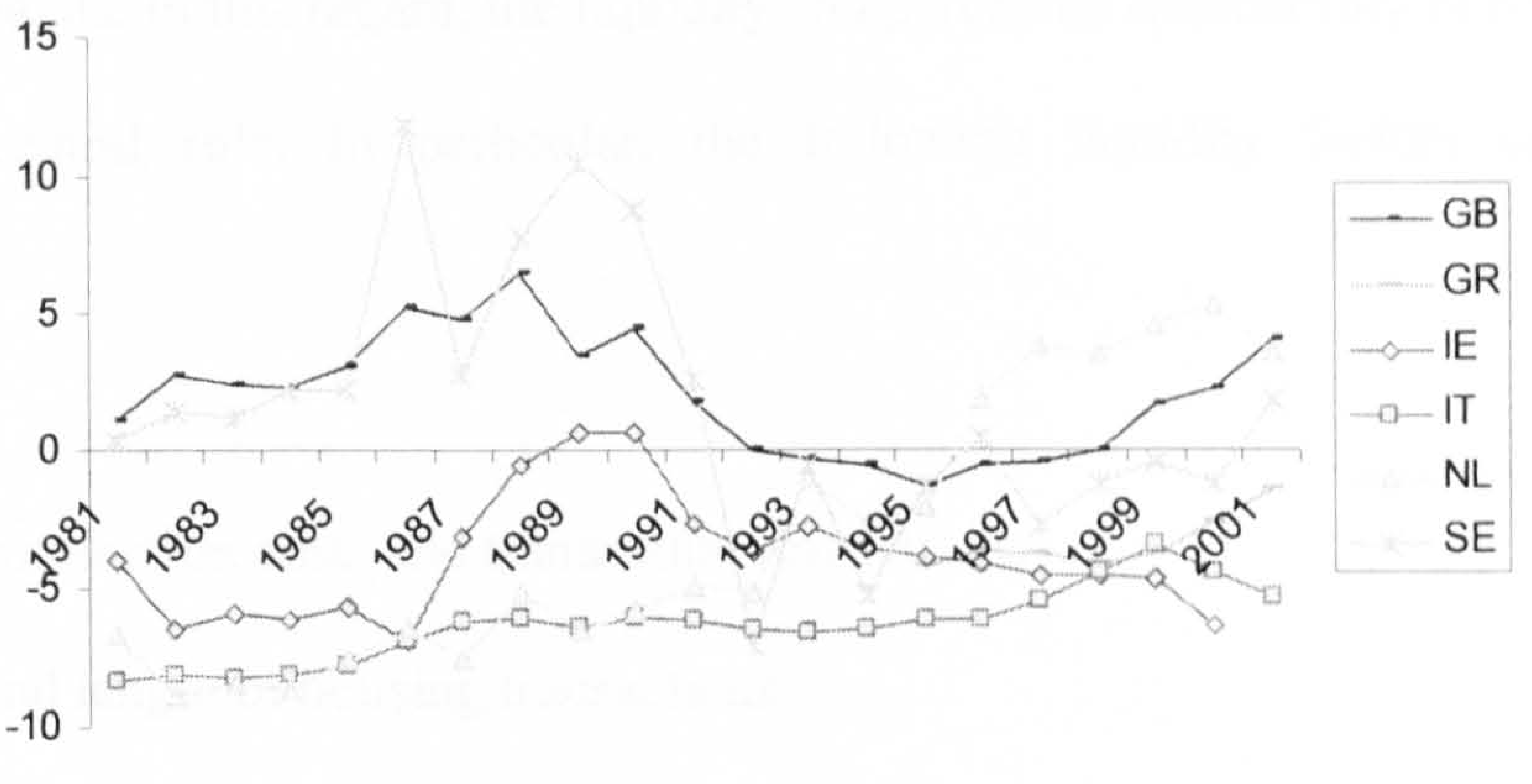
[16] In the U.K., Sweden, the Netherlands, Denmark and, to some extend (in the 1990s), Finland, it is possible to find a high and positive correlation coefficient between real house prices and the estimated MEW measure. This is consistent with an increased role of house prices as collateral in those countries with the most deregulated mortgage markets and easier access to housing lending. By contrast, where the deregulation of the mortgage market has been less intense, the household sector has been injecting equity into housing and the relationship between the two variables is less clear-cut.

FIGURE 2.2 Estimated Mortgage Equity Withdrawal (MEW) as Percentage of Disposable Income in the EU countries

Panel A



Panel B



Sources: Elaboration from ECB data. See Appendix in Chapter 3.
Notes: The MEW proxy is calculated as the change in the nominal mortgage debt minus nominal residential investment over household disposable income. AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, GR = Greece, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom. Portugal is not included.

The fifth and final channel which many economists have focused upon is the housing wealth effect to non-housing consumption expenditure. Miles (1994) argues that it is unclear why changes in house prices should be expected to affect private spending through wealth effects. The fact that demand increases must be matched by the supply from the selling agents (trading-down or last-time sellers) implies that, at the aggregate level, gainers and losers might tend to balance out their wealth or capital changes. Therefore, there is no *a priori* reason why an increase in real house prices should be treated as a boost in real wealth, unless houses are traded internationally.[17]

It is worth pointing out, however, that wealth effects could still be operating under specific conditions. Two possible explanations are: different marginal propensities to consume out of housing wealth (Muellbauer, 1995), and the presence of borrowing constraints. In this regard, the liquidity and perceived spendability of housing assets play a central role. In particular, the following liquidity factors can be identified:[18]

- housing transaction costs and transaction restrictions;
- number and length of housing transactions;
- collateral role of housing in the lending sector.

As far as the transaction costs are concerned, costs of real estate agents and lawyers, taxes and stamp duties are all restrictions on resale. They also reduce the probability of speculative frenzies by affecting and limiting the demand changes

[17] If a house in a country is bought by a foreigner, capital gains are enjoyed by the residents at the expense of capital loses of the buyers. This is what is happening in Spain and other Mediterranean countries where many British and North European citizens are buying second houses for retirement or holidays.
[18] These factors are also important in explaining the degree of sensitivity of housing demand and house prices to external shocks (e.g. mortgage interest rate movements, growth expectations, etc). Maclennan, *et*

deriving from interest rate movements. The number (as well as the average length) of transactions in the housing market is another factor of liquidity. This heavily depends upon the presence, as well as the efficiency, of real estate agencies in letting demand and supply for housing meet relatively quickly. Another factor is related to the collateral role of housing wealth and, in particular, the loan-to-value ratios, which reflect the characteristics of the credit market institutions as well as the legal and tax system. Maclennan *et al.* (1998) argue that, where housing is regarded as good collateral, housing wealth is more spendable and, therefore can affect the non-housing consumption to a greater extent. Some information on the above three liquidity factors will be provided in the next section.

Figure 2.3 summarises the above discussion, showing the main 'housing-system-related' channels of the monetary transmission mechanisms with their main inter-linkages. As in Figure 2.1, however, this simple diagram does not account for feedback effects and additional potential interactions between variables, which could be very relevant.

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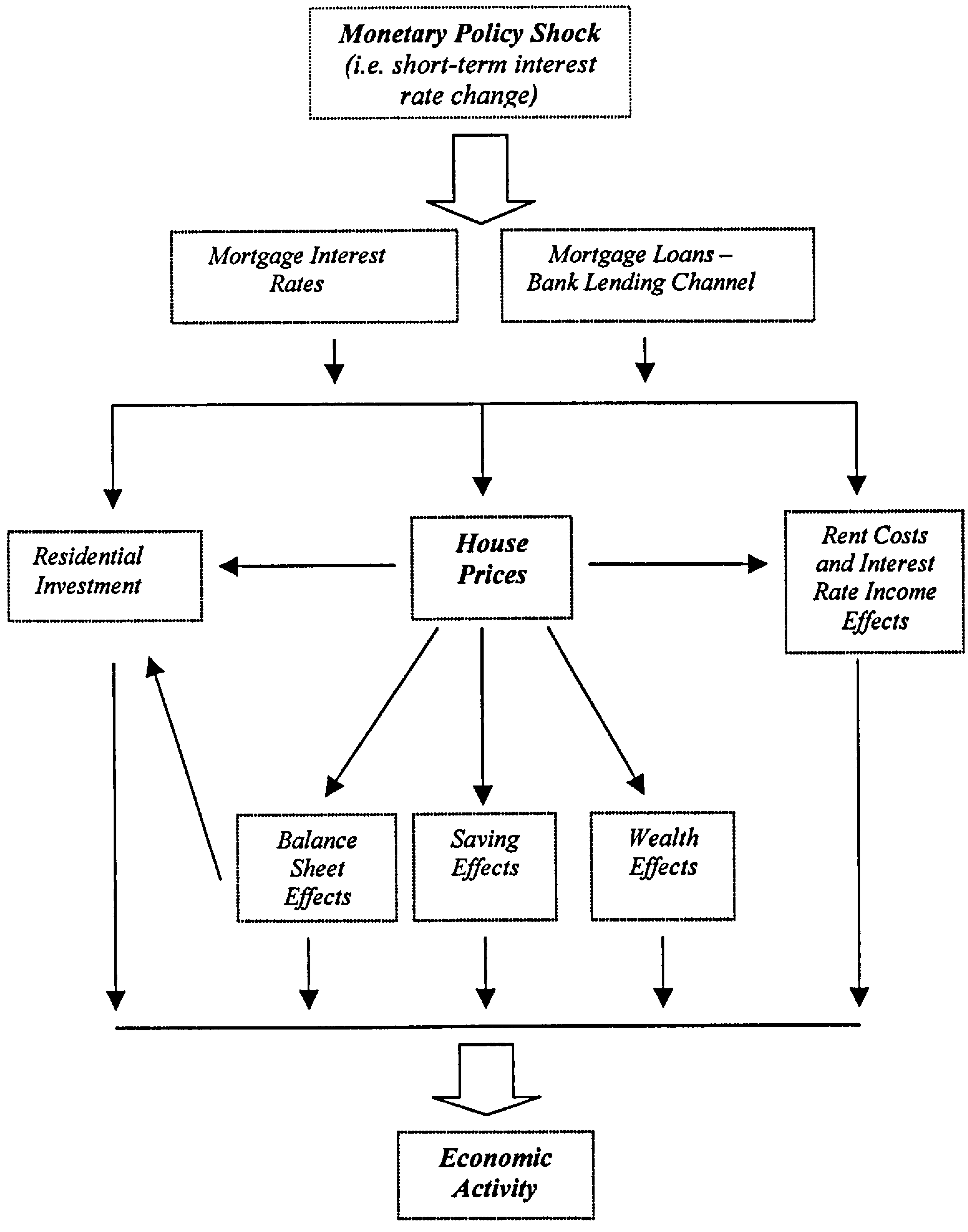
al. (1998) point to the role of asset price volatility, arguing that consumers are concerned about capital losses and the risk of loan default. In this respect, high volatility could reduce the spendability of housing.

FIGURE 2.3 The 'Housing-Related' Monetary Transmission Channels



Notes: For simplicity, this figure does not display all interactions between variables, but this can be important. Distributional effects across single individuals are not specified.

3. *Housing and Mortgage Systems across EU Countries*

The previous section has shown how several factors could affect the role of house prices in the transmission of a monetary shock into the households' spending and investment decisions. In particular, the following aspects have been emphasised: the size of the owner-occupied and private rented sectors, housing transaction costs, characteristics of the credit and mortgage markets and the mix between fixed or variable interest rate mortgages. In this regard, Maclennan *et al.* (1998) argue that:

"...countries with high (housing) transaction costs, low loan-to-value ratios, a small owner-occupied sector, a large tenure proportion in the private rented sector, and a large proportion of fixed interest mortgage loans, should experience relatively low real house price volatility, small house price effects on consumption and a small role for housing in interest rate transmission mechanism."

This section provides some evidence on the major housing institutional differences between the EU countries and tries to assess whether the housing sector (and in particular house prices) might play an important role in the MTM across European countries.[19] In analysing such features, it is important to point out that the comparability and availability of data represent one of the main weaknesses of cross-country studies, due to varying definitions and statistical sources.[20]

[19] The structure and the content of this section draws on the paper by Maclennan *et al.* (1998), which has become the main reference for cross-country studies of housing and mortgage systems. Recently, the ECB has published a report in which additional and more recent information has been collected (ECB, 2003).

[20] Appendix 1 provides some information in this respect. Additional information on the measures of mortgage market deregulation, house prices, mortgage stock and interest rate dynamics is offered in chapter 3.

Table 2.1 shows the overall tenure structure of the EU countries distinguishing between home-ownership, social housing and privately owned renting. This classification is particularly interesting, because a rise in real housing prices can be thought of as having positive wealth effects for owner-occupiers and negative income and substitution effects for tenants in the market rented sector. Therefore, the higher the percentage of owner-occupiers and the lower the proportion of households renting a house, the larger the house price effect will be on consumption decisions. Moreover, the collateral effects of house price movements on private spending will be stronger, the higher the proportion of owner-occupied.

Table 2.1 displays pronounced cross-country differences which could be explained by factors related to the heterogeneous fiscal treatment of home-ownership versus rental market, and historical and cultural reasons. As far as the importance of owner-occupied dwellings is concerned, Ireland, Spain, Italy, Greece and Luxembourg range between 72% and 80%, while Belgium, Finland, Portugal, Sweden and UK are in the range of 61% to 67%. All the remaining countries (namely, Austria, Germany, France, the Netherlands and Denmark) are particularly notable for owner-occupation rates which are well below the European average (63%).

Although the share of rented dwelling stock has decreased since 1980 in most European countries,[21] the proportion of the private rented sector is still very high in Portugal, Austria and Germany (respectively, 30%, 35% and 40%). Belgium, Italy and Greece range between 22% and 27% and all the remaining EU members have low private rental sectors (in particular Ireland and the United Kingdom). As emphasised by

[21] The main reasons being the strict rent controls and the increased demand for home ownership related to easier access to and lower cost to mortgage credit and favourable tax and subsidy policies. See Oxley and Smith (1996) and, more recently, the ECB (2003) for a more detailed discussion.

Maclennan *et al.* (1998), however, the role of the private rented sector (and the corresponding income effects linked to house prices movements) has to be moderated by the presence of rent controls, which limit the sensitivity of the rental market to house price dynamics and the overall efficiency of the housing system. The latter controls are even more dominant in 'social' rental accommodation, which represents more than half of the total rental sector in Finland, France, Ireland, the Netherlands, Denmark Sweden and the United Kingdom.

TABLE 2.1 Housing Tenure Structure in Europe

	Owner Occupation Rate	*Social Rental Occupation Rate*	*Private Rental Occupation Rate*	*Other*
Austria	50	15	35	0
Belgium	65	8	27	0
Finland	62	16	14	8
France	54	21	17	8
Germany	40	20	40	0
Greece	75	0	24	0
Ireland	80	10	10	0
Italy	75	3	22	0
Luxembourg	72	23	2	3
Netherlands	50	36	14	0
Portugal	66	4	30	0
Spain	78	0	18	4
Denmark	53	21	18	8
Sweden	61	22	17	0
United Kingdom	67	23	10	0
Euro Area	*64*	*13*	*21*	*2*
EU	*63*	*15*	*20*	*2*

Sources: Estimated from a variety of sources including the EMF (Hypostat, 1987-97, and Annual Report), Oxley and Smith (1996) and Maclennan *et al.* (1998).
Notes: Tenure expressed as % housing stock and refer to the mid-1990s. Data are not fully comparable for the presence of different definitions. See the above references for details.

The figures on the housing tenure can be better evaluated in relation to other important factors, which directly affect the sensitivity of housing demand and the 'perceived' liquidity or spendability of housing assets. In particular, the following aspects seem to be relevant: the presence of low transaction costs, easy credit availability, high loan-to-value ratios and a large diffusion of floating rate mortgages.

Unfortunately, as far as transaction costs are concerned, no data are available for all the EU countries. The figures provided by Maclennan *et al.* (1998) and ECB (2003) indicate that Spain and France present the highest proportion of total transaction costs, which on average respectively account for 10.4% and 13.8% of the house price. While the U.K. presents the lowest percentage in Europe (2%), Italy, Denmark and Germany (7.4% and 7.1%, respectively) are somewhere in the middle.[22] Others estimated figures published by the European Mortgage Federation (EMF, 2000) show that the total cost, including dwelling purchase (i.e. solicitor's fees or notary's fees, property registration and taxes) and mortgage loan costs (i.e. property valuation, mortgage registration, taxes on mortgages and solicitor's or notary fees) account for less than 3% of the purchase value in Denmark, Sweden and the United Kingdom and around 7-8% in France, Ireland, Italy, the Netherlands and Austria. In Greece it can reach up to 20%. Although not strictly homogeneous, the above figures show remarkable differences which are crucial not only for first-buyers but also home-movers.

An important element of the total transaction costs, whose quantitative relevance is easier to collect, are stamp duties. Data are not fully comparable in this case either and different sources provide different figures. Nevertheless, the ECB (2003) reports that Denmark, Germany, Finland, France, the U.K., Sweden and Portugal (between 1% and 4%) present the lowest rates. On the other side of the spectrum, there are Belgium

[22] These data refer to the first column of Table 5, page 20 in Maclennan *et al.* (1998) and the last column of Table 4.1, page 44 in ECB (2003). Data on total transactions costs (as well as its main components, i.e. stamp and registration duties, solicitors' fees, intermediation fees, etc) are extremely sketchy and differ from source to source.

and Greece (10-13%), and Italy, the Netherlands, Austria and Ireland which range between 6% and 9%. The features of housing taxes are further discussed in chapter 3, where additional information is provided on tax deductibility of mortgage payments and taxes on capital gains.

As emphasised in the previous section, another important factor affecting the sensitivity (and spendability) of house prices is given by non-monetary costs. For instance, in EMF (2000) it is shown that the typical length of time necessary to conclude a housing transaction (including the mortgage credit procedure) might take more than four months in Italy, around two in France and the Netherlands, and not more than a month in Denmark and Sweden. These figures are consistent with the number of housing transactions, which provides a measure of liquidity of the housing market. From EMF data, it emerges that over the period 1992-98, the (yearly) number of transactions per 1,000 residents was as follows: Italy, 8.7; Belgium, 10.3; Ireland, 13.5; Finland, 14.8; Denmark, 15.1; the Netherlands, 15.2; Sweden, 17.8 and the United Kingdom, 20.6.

Another influential feature of the housing lending market refers to the credit availability and, in particular, to the mortgage loans for owner-occupation. Despite the deregulation and liberalisation process of the 1980s and 1990s which affected the housing lending sector in most of the European countries, from Table 2.2 it is possible to see how there are still quite pronounced differences in the household mortgage debt to GDP ratio.[23] In particular, the latter ranges from 6-7% in Italy and Greece to 65% in Denmark. Sweden, the U.K. and Germany also present high ratios (between 50% and

[23] Data refers to 1998, which is the end year of our sample period. Chapter 3 provides a discussion of the housing lending debt to GDP ratio dynamics (Figures 3.2 and 3.3) and the main policy changes (Table 3.3) affecting the mortgage market in Europe during the 1980s and 1990s.

57%). Data on the number of new residential mortgage loans provides another indicator of credit ease. Over the period 1992-98 the number of new mortgages per 1,000 residents in Belgium, Ireland and Spain was respectively 13, 12.9 and 10.1. On the other side of the spectrum for the Netherlands, the U.K and Denmark, the respective figures were 26.2, 27.7 and 44.4.

Data on the typical loan-to-value (LTV) ratios available to borrowers are shown in column five of Table 2.2. With the only exception of Italy, which presents an extremely low 50% ratio, the remaining EU countries show ratios of around 70-80%, which are still well below the 90-95% ratio offered by the British banks. Although these figures refer to a specific year (namely, 1998) and are affected by the position in the country-specific housing cycle, they are particularly informative, because they provide an approximate measure of the down-payment requirements, and represent an indicator of the collateral role of housing. It is worth pointing out, however, that over the recent years the typical LTV ratios have shown a growing trend, although most EU countries have in place some mechanism rendering it costly to borrowers and lenders to agree on ratios above 75-80% (ECB, 2003).

Columns three and four of Table 2.2 provide some information on the contractual features of the existing mortgage stock, that is, the typical term maturity and the interest adjustments of mortgage debt. Although the financial deregulation process has increased the product availability making it very hard to define a typical contract in each country, Denmark is again an extreme case with a wide diffusion of mortgages with an interest rate fixed for the entire 25-year term. High levels of fixed interest rates can also be observed in France, Italy and Belgium. At the other end of the spectrum, there are Portugal, Finland and Spain (with, respectively, 100, 90 and 80% of variable interest rates), and the UK with almost all the mortgages categorised as renewable or reviewable. All the remaining countries are characterised by the presence of a mixture of fixed, variable, reviewable and renegotiable interest rates.

It is clear, however, that even under the presence of a similar mix of variable versus fixed interest rates, the size and the timing of the pass-through of policy rates to mortgage rates play a crucial role in the sensitivity of the demand for housing to monetary shocks and, under short-run rigid supply, of residential prices. In this regard, some authors (BIS, 1994; Borio and Flitz, 1995; Cottarelli and Kourelis, 1995; Mojon, 2000 and Toolsema, Sturm and de Haan, 2002) have examined the short-term and long-term pass-through of money market rates to various bank retail rates across a number of countries. Although results are not clear-cut and must be considered with caution due to the specific rates used, different empirical approaches and sample periods, they show that there is a significant heterogeneity, which can be partly explained by country-specific legal aspects of the banking sector (Cecchetti, 1999), the degree of competition and the integration of financial markets.

Looking at the effects of changes in money market rates on lending rates, the above studies show that the United Kingdom and the Netherlands have the fastest response, whereas Italy, Germany, Spain and Finland present a much lower degree of pass-through. As for the mortgage rates, BIS (1994) and Mojon (2000) show that Spain, Italy, Germany and France present a relatively slow short-term pass-through, whereas in the United Kingdom and the Netherlands it is found to be quicker. Additional evidence is given in Toolsema, Sturm and de Haan (2002) who, by focusing on Belgium, France, Germany, Italy, the Netherlands and Spain during the sample period 1980-2000, not only conclude that major differences in pass-through still exist, but that there is no clear indication of convergence of monetary policy transmission.

From the above analysis, it is difficult to draw a clear-cut identification of the countries where house prices could play an important propagating role in the MTM. Nevertheless, it might be useful to attempt a rough classification on the basis of the institutional features outlined so far. Table 2.3 provides the results of an 'experiment,' in which a number ranging between 1, 2 and 3 (indicating a relatively low, medium and high value) is assigned to the following five factors of the housing and mortgage markets:

- the ratio between the size of the rented market and ownership tenure;
- the number (and length) of housing transactions and low transaction costs;
- the mortgage lending availability and permissibility of products for MEW;
- the typical loan-to-value ratio;
- the mix between variable/fixed interest rate and the "pass-through".

Although it is difficult to assess the 'weight' or the relative importance of each factor, a visual inspection of the table suggests a relatively low potential role of house prices in the MTM in Belgium, Spain and Italy, while a relatively more significant amplifying role could be played in Sweden, Finland and the United Kingdom.

TABLE 2.2 Mortgage Systems in EU Countries

	Residential Mortgage Loans Stock as % GDP	*Interest Adjustment*	*Typical Term (years)*	*Typical LTV Ratio*	*Total Transaction Costs*
Austria	33	Some F Mostly N R	20-30	80%	7-8%
Belgium	22	N (40%) F (60%)	20	80%	Over 10-12.5%
Finland	30	V (90%)	10-15	70-80%	Over 4%
France	21	F (80%) V (20%)	15-20	70-80%	13.8% (1994) 7-8%
Germany	51	F (20%) N (40%) R (40%)	25-30	60-80%	7.1%
Greece	7	F (12%) R (72%) N (16%)	15	50-75%	Up to 20%
Ireland	27	F (64%) R (31%)	20	80%	7-8%
Italy	7	V (40%) F (60%)	10	50%	7.4%
Netherlands	60	V (10%) N (65%) F (25%)	30	75%	7-8%
Portugal	26	V (100%)	15	80%	> 10% in 1994
Spain	22	V (80%) F (20%)	15-20	70-80%	10.4%
Denmark	65	V (10%) F (90%)	25	80%	7.2% (2001)
Sweden	51	Mainly R and Short Term N	20-30	70-75%	Less than 3%
United Kingdom	57	R (70%) N (30%)	25	90-95%	Less than 3%
Euro Area	*27.8*		*19*	*73%*	
EU	*34.2*		*20*	*75%*	

Sources: EMF Annual Report (1998), Lea *et al.* (1997), Maclennan, *et al.* (1998) and Henley and Morley (2001), EMF (2000).
Notes: EMF data (1998) are mortgage loans for owner-occupation and rental purposes (not fully comparable). Fixed (F): rate fixed until final maturity; Renegotiable (N): rate not fixed over entire term, but more than one year; Variable (V) rate adjustable according to index; Re-viewable (R): rate adjustable at discretion of lender.

TABLE 2.3 Relative Potential Role of House Prices in the MTM

	Housing Tenure (Rented vs Ownership)	*Number (Length) of Transactions and Transaction Costs*	*Mortgage Availability and Equity Withdrawal (MEW)*	*Typical Loan-to-Value (LTV)*	*Fixed vs Variable Interest Rates and Pass-through*
Belgium	1	1	1	2	1
Finland	2	3	2	2	3
Ireland	3	2	2	2	1
Italy	2	2	1	1	1
Netherlands	2	2	3	2	2
Spain	2	1	1	2	2
Sweden	2	3	3	2	3
United Kingdom	3	3	3	3	3
Average	*2.1*	*2.1*	*2*	*2*	*2*

Notes: In all columns 1 indicates relatively low, 2 a medium and 3 a relatively high potential propagating role of the house prise in the monetary transmission mechanism. The average ratio between the size of the rented market and ownership tenure is 0.25. The only country above this number is Belgium (0.41). The U.K. and Ireland present a ratio equal to 0.15 and 0.12, respectively. The other countries range between 0.23 and 0.29. The second column summarises the (fragmented) information on number, length and cost of transactions discussed in the main text. Belgium and Spain have the lower number and/or the highest cost of housing transactions. On the other side of the spectrum, Finland, Sweden and the U.K. have relatively higher number of transactions (over 15 transactions per 1,000 people each year), lower costs (between 1-4%) and, when available, shorter transaction length (around one month). The third column summarises the information of the availability of housing lending and the permissibility of products for MEW. As for the former criterion, Belgium (22%), Italy (7%) and Spain (22%) are well below the average (34%). Although Finland (30%) and Ireland (27%) are still below this threshold, their markets have instruments facilitating MEW. As a result they have been classified with a 2. The remaining countries (the U.K., Sweden and the Netherlands) have ratios greater then 51%. The average typical LTV ratio is 70-80%. Italy is the only one below this value (50%), whereas the U.K. the only one above (90-95%). Finally, the last column provides a classification based on the diffusion of variable mortgage interest rates. Finland, Sweden and the U.K. have mainly variable interest rates (above 90%). Belgium, Ireland and Italy present shares of fixed rates above 60%.

The next sections try to quantitatively investigate the role of house prices for non-housing spending expenditure in the transmission of a monetary policy shock in two stages. Firstly, a study of the sensitivity of residential prices to interest rate policy changes using traditional unrestricted and semi-structural VAR models is carried out. The subsequent section assesses the size of those policy-induced residential price fluctuations for consumer spending decisions.

4. *The Empirical Methodology and Results*

The empirical method adopted in this chapter is motivated by two recent studies, which have tried to analyse how the impact of real asset values can magnify the effects of monetary tightening on the economy. In particular, Kwon (1998) extends existing work focusing on the transmission mechanism of monetary policy in Japan to a model with land prices, which play a primary collateral role in investment financing of Japanese firms. Following the standard empirical literature based on SVARs with restrictions on the contemporaneous interactions between variables, he extends Clarida and Gertler (1996) to a model with land prices and studies the effects of monetary policy on the variables of the system. After finding that a tightening of monetary policy leads to a significant and persistent decline in land prices and output, Kwon compares the decline in the latter variable to a model which does not include land prices. The main results show that the decline in output produced by a negative monetary shock appears more severe and deeper than estimated models without land prices predict. This finding is interpreted in favour of a significant propagating role of these assets in the MTM of Japan.

Another important contribution has been provided by Iacoviello (2000, 2002), who estimates a structural VAR for six EU countries (i.e. France, Germany, Italy, Spain, Sweden and the U.K.) using the common trends approach developed by King, Plosser, Stock and Watson (1991) (KPSW). In specifying a five dimensional VAR with output, real house prices, real money balances, short-term nominal interest rates and inflation as endogenous variables, Iacoviello shows that: (i) an adverse monetary shock has a significant negative impact on real house prices (with timing of the response in house prices matching that of output), (ii) the magnitude of the response in house prices to a monetary shock is partly explained by institutional differences in the housing and financial markets, and (iii) monetary and demand shocks affect house prices in the short run. In particular, by looking at the impulse response functions of the six countries to a standardised (50 basis points) short-term interest rate shock, Italy and UK present the largest fluctuations, with France and Germany at the other end of the spectrum. Spain and Sweden are somewhere in the middle. With the only exception of Italy, the results on the house price responses to monetary shocks are consistent with the institutional factors characterising the national housing and mortgage markets.

Although interesting, Iacoviello's analysis takes some short-cuts, which ought to be pointed out. First of all, the estimation period differs from country to country, making any quantitative as well qualitative comparison harder to assess. Moreover, it includes the post Bretton Woods flexible exchange rate regime and the semi-fixed or pegged exchange rate which followed the launch of ERM after 1979. As a result, not only his work might be distorted by the presence of multiple regimes, but his specification might not be rich enough, in that it omits variables (i.e. exchange rates), which are necessary to identify monetary policy shock in small open economies.

Given the main focus of this chapter (i.e. the investigation of the role of house prices in propagating the effects of monetary policy shocks), semi-structural VARs with restrictions on the contemporaneous interactions between variables are used. This approach has a number of advantages as well as problems, which are discussed below. Before showing the main results and motivating the specific identifying restrictions, the next section briefly outlines the structural VAR methodology.

4.1 *The Structural VAR Approach*

Letting $\mathbf{x}_t$ denote the *(nx1)* vector of endogenous variables and $\boldsymbol{\varepsilon}_t$ the *(nx1)* vector of structural shocks, it is assumed that the economy is represented by a linear, stochastic dynamic system of the following structural form:

(1) $$\mathbf{B}_0\mathbf{x}_t = \mathbf{B}_1\mathbf{x}_{t-1} + ... + \mathbf{B}_p\mathbf{x}_{t-p} + \boldsymbol{\varepsilon}_t$$

where $\mathbf{B}_i$, $i = 0,...n$, are *(nxn)* matrix of coefficients and where the structural disturbances $\boldsymbol{\varepsilon}_t$ are assumed to be mutually orthogonal and have unit variance, implying that $E(\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t^{'}) = \mathbf{I}$.

The moving average (MA) representation for $\mathbf{x}_t$ is given by:

(2) $$\mathbf{x}_t = [\mathbf{B}(L)]^{-1}\boldsymbol{\varepsilon}_t$$

where L is the lag operator (that is, $L^k\mathbf{x}_t = \mathbf{x}_{t-k}$) and $\mathbf{B}(L) = \mathbf{B}_0 - \mathbf{B}_1L - \mathbf{B}_2L^2 - ...\mathbf{B}_pL^p$. Eq. (1) can be written in a reduced form, which can be directly estimated by ordinary least square (OLS), as:

(3) $$\mathbf{x}_t = \mathbf{A}_1\mathbf{x}_{t-1} + ... + \mathbf{A}_p\mathbf{x}_{t-p} + \mathbf{u}_t$$

where the innovations $\mathbf{u}_t$ are assumed to have variance-covariance matrix $E(\mathbf{u}_t\mathbf{u}_t^{'}) = \boldsymbol{\Sigma}$.

Noting that $\mathbf{A}(0) = \mathbf{B}_0^{-1}$, it is possible to show that $\mathbf{A}_i = \mathbf{A}(0)\mathbf{B}_i$, for $i = 1,...n$, and the relationship between the innovations $\mathbf{u}_t$ and the structural shocks $\boldsymbol{\varepsilon}_t$ as:

(4) $$\mathbf{u}_t = \mathbf{A}(0)\boldsymbol{\varepsilon}_t \quad \text{or} \quad \boldsymbol{\varepsilon}_t = \mathbf{B}_0\,\mathbf{u}_t$$

from which it can be easily seen that:

(5) $$E(\mathbf{u}_t\mathbf{u}_t^{'}) = \boldsymbol{\Sigma} = \mathbf{A}(0)\mathbf{A}(0)'$$

From eq. (2), it is straightforward to calculate the impulse response functions (IRFs) to the structural shocks ε_t as $\mathbf{x}_t = [\mathbf{B}(L)]^{-1}\varepsilon_t = \Phi(L)\varepsilon_t$ where from eq. (3) and eq. (4):

(6) $\quad \Phi(L) = [\mathbf{I} - \mathrm{A}(L)]^{-1}\mathrm{A}(0)$.

While OLS can be used to obtain consistent estimates of the parameters in (3) and an estimate of the symmetric variance-covariance matrix Σ, to identify the structural shocks ε_t and the above IRFs coefficients it is necessary to determine the n^2 elements of $\mathrm{A}(0)$. As is well known in the literature, system (5) provides n^2 equations in only *n(n+1)/2* restrictions defined by Σ. Just-identification of such a model and, therefore, estimation of the matrix $\mathrm{A}(0)$ and the structural innovations ε_t require *n(n-1)/2* additional restrictions.[24]

As comprehensively explained in Favero (2001), the most common strategies used by the VAR literature imply the imposition of restrictions (i) on the contemporaneous effects of reduced-form shocks $\mathrm{A}(0)$ (Sims, 1980 and 1992); (ii) on the contemporaneous effects of the endogenous variables $\mathbf{B}_0$ (Bernanke, 1986 and Sims, 1986); (iii) long-run *a priori* theoretical restrictions on $\mathbf{B}(1)$ or $\mathrm{A}(1)$ (Blanchard and Quah, 1989; Clarida and Galí, 1994); and some combination of (i), (ii) and (iii) (Christiano *et al.*, 1999 and Galí, 1992).

A standard way to add the required additional restrictions is to use a Cholesky decomposition of the variance-covariance matrix Σ and, therefore, assume that $\mathrm{A}(0)$ is lower triangular. This assumption implies a recursive economic structure, in which a

[24] It is correct to say that exact identification requires not only this order condition, but also a rank condition. See Hamilton (1994) for a discussion.

shock on variable i contemporaneously affects all the variables j if, and only if, $j \geq i$. In the following sections, both recursive and semi-recursive (semi-structural) restrictions are imposed.

4.2 *The Role of Monetary Shocks for House Price Fluctuations*

For a real asset price channel to exist, two conditions have to be satisfied. First, significant effects in the house prices should accompany shocks in the interest rates. Second, these fluctuations should affect the economic activity mainly through changes in the non-housing consumption of households.[25]

This section starts by estimating a recursive VAR model to focus on the first condition and assesses the role of monetary policy shocks for housing price developments. Due to data availability, the eight countries under investigation are Belgium, Finland, Ireland, Italy, the Netherlands, Spain, Sweden and the U.K..[26] The estimation period is 1979:2-1998:4 for all countries, with the exception of Spain (1987:1-1998:4) and Belgium (1981:1-1998:4). Our specification includes, in the following order: the consumer price index (CPI), the real gross domestic product (GDPV), the real house prices (RHP) and the money market interest rates (IRS).[27] Given the short-run focus of the analysis, it is assumed that the stock of houses is fixed

[25] It is clear that the economic activity is also affected through residential investment changes. See Bernanke and Gertler (1995) and, more recently, McCarthy and Peach (2002) for a study of the effects of monetary policy on residential investment in the United States.
[26] Differently from Iacoviello (2000), France and Germany are not included in the sample. While for the latter country only annual data are available, in the case of France, the quarterly index refers to the Paris area only and cannot be considered a good representation of national house price dynamics. See Appendix 1 for a full description of the time series used in this chapter and their sources.
[27] As in Sims (1986 and 1992) and the subsequent VAR literature, innovations in short-term interest rates are used as a measure of shocks to monetary policy. Although central banks have a wider set of policy instruments at their disposal, in achieving a common specification of instruments, which might help to achieve a minimum degree of comparability, the assumption that during the sample period 1979-98, the individual European CBs have used short-term interest rates as their main tool does seem more than reasonable (Borio, 1997).

and housing supply variables are not included. This is not a particularly important omission, because house supply in Europe has been shown to be very inelastic in the short term (Ball and Grilli, 1997). Additionally, it is not possible to obtain quarterly series of housing stock for the vast majority of the countries under investigation.

In imposing the Cholesky decomposition with the above ordering, it is assumed that monetary policy actions do not have any contemporaneous effect on the economic activity and that authorities are able to take into account the simultaneous developments of the economic variables in setting their policy.[28] All variables are taken in natural logarithms, with the exception of the money market rates, which are in levels.

Although some of the variables might have a unit root, all the models are estimated in levels.[29] This choice is made for a number of reasons. Firstly, if cointegration among the variables exists, the system's dynamics can be consistently estimated in a VAR in levels (Sims, Stock and Watson, 1990). Given the loss of efficiency associated with the latter, an alternative approach is to use a VAR in first differences alone. This solution, however, implies discarding the information contained in the levels, and could lead to mis-specification. An intermediate method would be to estimate a Vector Error Correction Mechanism (VECM), but imposing inappropriate cointegrating relationships may seriously bias the impulse response (Hamilton, 1994

[28] Alternatively, policy variables can be ordered first, which implies assuming that monetary shocks can contemporaneously affect the other variables of the system and that monetary authorities do not consider the simultaneous innovations of the economic variables (Sims, 1992). While this identification assumption might be reasonable for high frequency (i.e. monthly) data, the chosen ordering seems to be more appropriate with quarterly time series. However, different orderings of the four variables have been tested. Due to the low correlation between the residuals of the reduced-form equations (Enders, 1995) results are not too much affected.

[29] Standard unit root (Augmented Dickey Fuller and Phillip-Perron) tests were implemented. Most of the variables used in the VAR models below are integrated of order one. The only exception is the level of the consumer price index, which, in some countries, results stationary around a deterministic trend, whereas in others appears to be integrated of order one or two. Given these mixed results and the low power of the tests, the robustness of the estimated models has been tested using both the (log)-level and the growth of the CPI. The empirical findings are consistent across the two specifications, and the impulse responses of the key variables to monetary shocks not too much affected.

and Hendry, 1996 for a discussion). On the basis of these considerations and the lack of *a priory* cointegrating relationships to impose on the data, similarly to a number of empirical studies, the VAR systems below are estimated in levels.[30]

The benchmark VAR system is estimated with different lag lengths according to the selection made by the Akaike Information Criterion (AIC) and the Schwartz Bayesian Criterion (SC).[31] Longer lag lengths do not change the main qualitative results, although the rapid loss of degrees of freedom affects the statistical significance of some impulse response functions. The residuals are however checked for serial correlation up to the fourth order, and all the tests are passed at standard level of significance. In order to avoid heteroskedasticity problems and increase efficiency of the estimated parameters, a number of dummies have been included for Ireland, Italy, the Netherlands and Sweden to take outliers into account, which were mainly observed in the reduced-form interest rate equations during the exchange rate crisis of 1992-93. A set of seasonal dummies is also included.

Figure A.2.1 in Appendix 2 shows the point estimates (and the corresponding two-standard error bands are constructed using bootstrapping techniques with 1,000 replications) for the responses of the CPI, real GDP, RHP and IRS to an exogenous and temporary monetary policy shock. The restrictive impulse generates a significant and temporary decline in output for all countries. The only exception is Spain, where the real output fall is relatively small and not statistically significant. Once the size of the shock to the money market rate is taken into account, differences in the timing and size

[30] Unrestricted VARs in levels with alternative contemporaneous identifying restrictions have been applied by Sims (1986, 1992), Bernanke (1986), Kim (1999, 2000, 2002), Christiano, *et al.* (1999), Dedola and Lippi (2000), Peersman and Smets (2001), Clements, *et al.* (2001) amongst others.

[31] In particular two lags were selected in all countries, with the exception of Ireland and Belgium (three lags) and Italy (four lags).

are very noticeable and similar to previous empirical studies.[32] In the Netherlands and Finland, the increase in the short-term interest rate results in a temporary (but quite persistent) increase of the price level which, however, permanently decreases after 15-20 quarters.

This 'price puzzle' has been found in several empirical studies and, given the low number of endogenous variables included in the model, it could be due to a mis-specification of the system, where leading indicators for inflation, to which central banks react, might be omitted (Sims, 1992).[33] The main successful solution proposed by the empirical U.S. literature relies on adding current and lagged values of commodity or oil prices into the monetary authorities reaction function to take into account supply effects and ordering the short-term interest rate after these variables (Christiano, *et al.*, 1999). As in Sims (1992) and a number of other studies, results are not particularly satisfactory for European countries. In fact, the point estimates of the response (results not shown) of the price index remain unaffected when a commodity price index is included.

As far as the real house price effects are concerned, it is possible to identify three main groups of countries, roughly consistent with the descriptive analysis of the housing and mortgage systems across Europe. Although a rigorous comparison of the point estimates is not warranted by the presence of heterogeneous definitions and construction (see Appendix 1 for a discussion), in the first group there are Spain, Belgium and Italy. While in the former two countries, a small and not statistically

[32] See Peersman and Smets (2001) for a recent survey on the MTM studies based on the VAR approach.
[33] As argued in Favero (2002), if such a leading indicator for inflation is omitted from the VAR, we have an omitted variable positively correlated with inflation and interest rates, explaining the positive relation between the latter two variables found in the impulse response functions. Also by adding the relevant variables, it might be the case that monetary authorities do not react with sufficient commitment to offset future inflationary pressures, because of the necessity of smoothing changes of interest rates over time.

significant temporary decrease of the real house price index is found, in Italy it is possible to observe an anomalous response, with an increase of the residential price index over the first 8-10 quarters, which quickly reverses to a significant but temporary medium-run decrease.

On the other side of the spectrum, Finland, Sweden and the U.K. show the most significant and largest decline. Ireland and the Netherlands are somewhere in the middle. Similarly to Iacoviello (2000), the timing and the responses in house prices matches that of real output, providing some evidence of the fact that the effects of monetary shocks to house prices work through changes in the real activity. In order to enhance comparability of the IRFs, Figure A.2.1a in Appendix 2 graphs the point estimates of the house price responses to a normalised (100-basis-points) shock to the interest rate. From this diagram it can easily be verified that the countries under investigation show a heterogeneous sensitivity of residential prices to monetary shocks. In particular, a temporary 100-increase of the short-term rate generates a U-shaped fall of the real house prices in most countries. The maximum effects (between 2.5-3%) can be easily seen in Finland, Sweden and the United Kingdom.

The above specification describes a closed economy model. This might not be an ideal choice if the aim is to identify unexpected monetary shocks for relative small and open economies. Moreover, the estimation period has been characterised by exchange rate targeting and management of the currency parities. As a result, the robustness of the above four-variable system is tested by augmented it with the nominal exchange rate versus the DM (NER). Under the Cholesky decomposition, however, it is hard to find an ordering which can be unquestionable. Ordering the exchange rate before the interest rate would imply the assumption that the latter does not contemporaneously affect the former. The reverse would impose no current feedback reaction of the monetary authorities to exchange rate fluctuations over the same quarter. Bearing this in mind, both orderings have been imposed with no evidence of significant differences in the estimated impulse response functions. In most countries, however, it is found that a rise in home short-term interest rates is associated with contemporaneous domestic exchange rate depreciations (i.e. exchange rate puzzle). This implies that the Cholesky decomposition is not advisable in VAR specification, where both interest rates and exchange rates are included. Therefore, the basic closed-economy specification is kept as the benchmark system, testing for alternative decompositions below.

A sub-sample stability analysis to verify the robustness of the above results is carried out. In particular, given the limited number of observations, the models are re-estimated over the sample period 1985:1-1998:4. There are two reasons for the choice of this sample. First, it is possible to check the stability of our point estimates under the so-called "hard" EMS regime (Ehrmann, 2000). Secondly, this sample allows us to verify the importance of short-term interest rate shocks during the period of major expansion of mortgage credit liberalisation process, which started in the early 1980s but, in many countries, developed fully during the second half of that decade.[34] The system has also been estimated by adding a proxy for the credit availability (e.g. total credit to the private sector) over this sample. In both cases, results are to a large extend unaffected and the simulated impulse response functions show similar dynamics. This suggests that, although rather simple, the recursive structure represents a robust specification.

[34] Iacoviello and Minetti (2002) study the sensitivity of the house prices to monetary policy in Finland, Sweden and the U.K. under different financial liberalisation regimes and choose a similar sub-sample period.

While recursive identifications schemes with an appropriate choice of endogenous variables have been shown to provide satisfactory results for relatively closed economies like the U.S. (Christiano *et al.*, 1999), as shown and verified above, the Cholesky decomposition has been less successful in studying the MTM in small open economies with the inclusion of exchange rates. Following Sims (1986), Bernanke (1986), and more recently Kim (2001, 2002) and Kim and Roubini (2000), a semi-structural VAR with identification based on contemporaneous exclusion restrictions on the coefficient matrix $\mathbf{B}_0$ (see eq. (4) in section 4.1) is implemented. This method is particularly appealing, because it explicitly allows for the simultaneous estimation of the interactions among variables within the same quarter.[35]

In particular, in each country under investigation the previous specification is augmented with the nominal exchange rates versus the DM (NER).[36] While this choice clearly satisfies the necessary requirements of an identical structure, it does not reflect actual differences in the CBs reaction functions.[37] Although some heterogeneity in the specifications is allowed by using different lag lengths and including appropriate dummies for special events, this is costly in that some puzzling responses to monetary policy shocks still remain.[38] At the same time, however, it provides a common specification which enhances the comparability of the estimation results.

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[35] In the Cholesky decomposition, by including some variables in the monetary reaction function, it is assumed that monetary policy shocks do not affect those variables contemporaneously, and vice versa. While this assumption is not particularly costly with economic variables, the inclusion of nominal exchange rates in the specification model makes the recursive identification theoretically weak.

[36] Following Ehrmann (2000) for Ireland and Finland the Punt/Sterling and Markka/US$ exchange rates are used. While from a theoretical point of view this is justified by monetary authorities, who, in the Irish case, might have been more directly affected by those rates, in Finland this choice prevents any puzzling responses to monetary shocks.

[37] For instance, to identify a monetary policy disturbance in many of the countries under investigation (in particular, the Netherlands and Belgium), it might be more accurate to include German interest rates and output as additional exogenous variables (see Peersam and Mojon, 2002 for a similar specification choice). This solution, however, implies a rapid loss of degrees of freedom which makes the results very unstable.

[38] The construction of country-specific models could be an interesting extension.

The estimated system is given by the following variables: the consumer price index, the real output, short-term rates, real house prices and nominal exchange rates. Following Sims (1986) and Bernanke (1986), in order to achieve identification of the structural parameters and innovations, a set of zero (exclusion) restrictions on the contemporaneous structural parameters $\mathbf{B}_0$ are imposed. They can be summarised in the following system:

$$\begin{bmatrix} \varepsilon_{cpi} \\ \varepsilon_{gdp} \\ \varepsilon_{ms} \\ \varepsilon_{rhp} \\ \varepsilon_{ner} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & a_{35} \\ a_{41} & a_{42} & a_{43} & 1 & 0 \\ a_{51} & a_{52} & a_{53} & 0 & 1 \end{bmatrix} \begin{bmatrix} u_{cpi} \\ u_{gdp} \\ u_{irs} \\ u_{rhp} \\ u_{ner} \end{bmatrix} \tag{7}$$

where ε_{cpi}, ε_{gdp}, ε_{ms}, ε_{rhp} and ε_{ner} are the structural disturbances, that is, CPI, GDP, monetary, house price and exchange rate shocks,[39] respectively, and, u_{cpi}, u_{gdp}, u_{irs}, u_{rhp} and u_{ner} are the innovations of the reduced-form equations.

This structural decomposition differs from a standard Cholesky factorisation in two equations. The interest rate equation is assumed to be a reaction function of the monetary authorities, who, within the quarter, can observe the current value of the price and output level, as well as the nominal exchange rate. The nominal exchange rate equation is assumed to be dependent on all of the variables in the system except for the real house price index, which, apart from being publicly available with delays, might be expected to have a secondary role in the exchange rate determination. The remaining zero contemporaneous restrictions in the other equations are consistent with a recursive

[39] Given that this is a model geared to provide monetary policy shocks, this chapter does not consider alternative restrictions, which might help to better identify (and interpret) the other structural shocks of the system.

ordering given by {CPI, GDPV, IRS, RHP, NER} and have been widely used by the relevant literature.

The system is just-identified and the contemporaneous coefficient matrix $\mathbf{B}_0$ has been estimated by maximum likelihood estimation available in RATS. Similarly to the benchmark model, all variables enter in logarithms, except for short-term interest rates. A complete set of seasonal dummies is used in all estimations and each variable enters each equation with two or three lags, according to AIC and SC. Higher orders have been tested and do not affect the main results.

Figure A.2.2 reports the IRFs to the identified monetary policy shock. A temporary price puzzle still occurs in some countries, but, in all cases, it is not statistically significant and much less persistent than in the recursive structure.[40] The short-run appreciation of nominal exchange rate in most countries provides some positive support in favour of a successful identification of the monetary policy shock. As far as the real output responses, the imposed structure seems to match the recursive system results. In the case of Finland, however, the effects of monetary policy on real house prices are reduced in size. Similarly to above, the same model is re-estimated during the so-called "hard" ERM. The main qualitative results seem to be very robust, although the loss of degrees of freedom makes the estimated responses less stable.

At this point, it is interesting to study the variance decomposition of the real house price index, and to verify which weight can be attributed to the different innovations of the system and, in particular, to the identified monetary policy shock. Figure A.2.3 shows the share of the total variation of the real house price attributable to

[40] With the exception of Spain, the CPI response becomes permanently negative 10 quarters after the monetary shock.

each specific shock. It is worth pointing out, however, that the SVAR models, as the ones estimated in this section are mainly designed to identify monetary shocks and, as a result, it is not possible to label or interpret the remaining shocks.[41]

The figure indicates that in all countries, residential price volatility is mainly driven by innovations in the real house price equation. However, consistent with the above IRF analysis, the Netherlands, Sweden, the United Kingdom and, to some extent, Finland are the countries in which monetary shocks play a relatively more significant role (after the 4-5 quarters it goes from 5% to 42%). On the other side of the spectrum, there are Spain, Italy and Belgium, where the role of interest rate shocks is marginal or very small. Shocks to the real GDP equation are particularly relevant in driving house price fluctuations (between 20-40%) at any time horizon in Belgium, Ireland, Spain, Finland and Italy. Similarly to Iacoviello (2000), the identified monetary shocks seem to play an influential role in the housing market. However, the estimates above provide some evidence on the possibility of cross-country differences, which are more closely related to the institutional features of the European housing and mortgage markets outlined in section 3.

4.3 The Role of House Prices in the MTM to Consumption

So far, attention has been focused on how exogenous interest rate shocks affect real house prices. However, the discussion provided in section 2 has emphasised the potential role of house price fluctuations in affecting non-housing consumption expenditure of households. The purpose of this sub-section is to investigate the quantitative significance or amplifying role of these effects in the transmission of

[41] A different identification scheme with some combination of long and short-run restrictions would have been more appropriate to identify the other structural shocks (Gall, 1992; Clarida and Gall, 1994).

interest rate shocks to non-housing consumer spending in a SVAR system. In order to achieve this, two approaches are implemented. The first is based on a counterfactual simulation exercise, the second on the estimation and comparison of models in which the key variable (e.g. house prices) enter the systems endogenously and exogenously.

4.3.1 A Counterfactual Simulation Exercise

Following Kim (1995), Sims and Zha (1995), Bernanke, Gertler and Watson (1997), and more recently, Lettau, *et al.* (2002), this sub-section implements a two-stage approach in a VAR context. In the first stage a 'baseline' structural model which includes real output, consumer price index, real private consumption, short-term rates, real house prices and exchange rates is estimated. Similarly to the semi-recursive system of the previous section, the exogenous structural monetary shock is identified through restrictions on contemporaneous relationships between the reduced form and the structural innovations, which can be summarised by the following system:

$$(8) \qquad \begin{bmatrix} \varepsilon_{cpi} \\ \varepsilon_{gdp} \\ \varepsilon_{cons} \\ \varepsilon_{ms} \\ \varepsilon_{rhp} \\ \varepsilon_{ner} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 & 0 \\ a_{41} & a_{42} & a_{43} & 1 & 0 & a_{46} \\ a_{51} & a_{52} & a_{53} & a_{54} & 1 & 0 \\ a_{61} & a_{62} & a_{63} & a_{64} & 0 & 1 \end{bmatrix} \begin{bmatrix} u_{cpi} \\ u_{gdp} \\ u_{cons} \\ u_{irs} \\ u_{rhp} \\ u_{ner} \end{bmatrix}$$

where ε_{cpi}, ε_{gdp}, ε_{cons}, ε_{ms}, ε_{rhp} and ε_{ner} are the structural disturbances, that is, CPI, GDP, real private consumption, monetary policy, house price and exchange rate shocks, respectively, and, u_{cpi}, u_{gdp}, u_{cons}, u_{irs}, u_{rhp} and u_{ner} are the innovations of the reduced-form equations.

This structural decomposition is qualitatively similar to the one used in the previous section, with the only difference that the reaction function of the monetary authorities is augmented with the current and lagged values of the real private consumption. Again, the nominal exchange rate equation is assumed to be dependent upon all variables in the system except for the real house price index. The remaining zero contemporaneous restrictions in the other equations are consistent with a standard Cholesky ordering implying that real private consumption is not contemporaneously affected by interest rates and house price fluctuations. The latter assumption is suggested by the fact that the actual house price dinamics might not be readily available within the same quarter and, as a result, consumer spending is assumed not to be affected by contemporaneous changes in real asset values. Alternatively, this restriction can be thought of as a sluggish response from consumers in their spending decisions.

The above system is used to estimate IRFs, which provide the total dynamic responses of private non-housing spending to the identified interest rate shock. It is worth pointing out that, in excluding the contemporaneous effect of residential prices to private consumption, the operativeness of the house price channel is not shut off. In fact, from the structural system provided by equation (1), it can be easily seen that lagged house prices can still directly affect consumers' decisions through the channels emphasised above.

In the second stage, the effects of such shocks are simulated under a counterfactual regime in which the effects of house price fluctuations to private consumption are 'shut-off.' This is done by setting to zero any response of consumption to house prices and keeping the remaining structural parameters fixed to their baseline value. The effects of money market rate innovations on private spending are then re-computed and compared with the ones estimated in the first stage. Similarly to Lettau *et al.* (2002), the difference between the two responses can be interpreted as a measure of the propagating contribution of the house price in the effects of a monetary shock to household consumption.

Figure A.2.4 shows the responses to a monetary shock in the six-variable system.[42] As for the variables common to the previous models, the quantitative and qualitative responses are basically unchanged and support the main findings on the sensitivity of the house prises to the identified monetary policy shock.[43] Additionally, a statistically significant U-shaped response of private consumption is found. The counterfactual simulation results are shown in Figure A.2.5. From a visual inspection, it is interesting to see that the propagating role of house price on consumer spending is particularly strong in the U.K., Finland, the Netherlands and Sweden, and marginal in the remaining countries, where the two responses almost coincide.

4.3.2 Endogenous versus Exogenous House Prices Models

To test for the sensitivity and robustness of these above results, an alternative approach aiming to disentangle and quantify the strength of the house price channel is here implemented. In particular, following Morsink and Bayoumi (1999) and Clements *et al.* (2001), a recursive VAR in levels (with the following variables: consumer price index, real GDP, real private consumption, real house prices and short term interest rates) is estimated.[44] This model is then re-estimated, eliminating the house price equation while maintaining the lagged values of the real house price index in each equation of the

[42] Due to convergence problems with the six-variable system for Ireland, results of the counterfactual experiment are not reported.
[43] The only exceptions are Spain and Italy with the presence of an exchange-rate puzzle.
[44] In this sub-section a closed-economy specification is used for two main reasons. Firstly, from a preliminary analysis, while that inclusion of the exchange rates provides a more reliable identification of the monetary shock in the pre-EMU period, this leaves the quantitative responses of real house prices and consumption unaffected. Secondly, by using a different specification and a non-recursive approach, it is possible to test for the robustness of the counterfactual simulation results. The sensitivity of the results to a number of alternative specifications and different lag lengths structures have been tested, without finding significant changes.

system. In doing so, the latter model (in which residential prices are treated as exogenous) blocks off any effect within the VAR that pass-through the endogenous house price response to the interest rate shock.[45] Similar to the counterfactual experiment shown above, comparing the responses of the two models to monetary policy shocks provides a measure of the role of the house price channel in the MTM.

Figure A.2.6 shows the results from which it is possible to confirm the conclusions of the counterfactual simulation.[46] In particular, the U.K., Finland, the Netherlands and Sweden are the countries where house prices seem to play amplifying role in the MTM to private consumption. Namely, in the former two countries the household's expenditure response to interest rate shocks is circa halved, whereas in the Netherlands and Sweden the effects are reduced by around one third. On the other hand, in Belgium and Italy the two responses are very similar.

In general, the results shown in Figure A.2.5 and A.2.6 are consistent with the existence of a positive housing wealth and collateral effects in those countries where a greater sensitivity of house prices to monetary policy changes was found. However, it is worth pointing out that also in the extreme cases, the estimated IRFs in the two exercises lie within the two-standard-error confidence bands of the respective baseline models. While this does not invalidate the main qualitative findings, wide confidence intervals which are an inevitable price to pay in over-parameterised systems such as the ones estimated above. In this respect, an alternative estimation approach characterised

[45] A block exogeneity (or 'block exclusion') test was carried out to determine whether lags of the real house price index enter the equations for the CPI, real GDP, real private consumption and short term interest rate equations (see Enders, 1995, p. 316 for an outline). The likelihood ratio statistic (Sims, 1980, p. 17) rejects the null hypothesis of block exogeneity at 5% level in the U.K., Finland and the Netherlands. For the remaining countries (namely, Sweden, Italy, Belgium and Spain) the tests cannot reject the exogeneity hypothesis at any standard level of significance.

[46] Irish and Spanish estimates did not provide reliable and stable impulse responses for most of the variables. Therefore, results are not reported.

by more parsimonious specifications might strengthen the statistical significance of these results. The following chapter will explore this aspect further.

5. Conclusions

Over the last few years many researchers have been trying to compare and analyse the effects of monetary policy shocks across European countries. While the evidence is inconclusive and rather mixed, few of them focus on single channels of the MTM. Even fewer have attempted to assess the role of house prices. This is rather surprising, because housing represents more than half of the net wealth of the private sector, and plays a crucial role as collateral for households' borrowing. Moreover, the last decades have witnessed large residential price fluctuations, which have been shown to have affected the real economy in many industrialised countries (Kennedy and Andersen, 1994; Muellbauer and Murphy, 1997; Girouard and Blöndal, 2001; Ludwing and Sloek, 2002 and Case *et al.,* 2002). In this respect, while there is no controversy on the importance of residential house prices on the real economy, little work has been done in investigating the role of such assets in the transmission of interest rate shocks.

This chapter provides some qualitative and quantitative evidence of the link between monetary shocks and residential prices, and between house prices and economic activity over the EMS period for eight European countries, by using alternative structural VAR models. From the estimated impulse response functions and variance decomposition analyses, house prices are found to be significantly affected by monetary shocks in most of the countries under investigation. In particular, the results show that temporary restrictive unexpected monetary policies have short-run negative effects on real house prices and that the size of the response is partly consistent with differences in the housing and credit systems. In particular, in those countries where mortgage markets are relatively more competitive and housing systems more efficient, like Finland, Sweden and the United Kingdom, residential prices are more sensitive to interest rate shocks. On the contrary, in countries like Italy, Spain and Belgium this does not seem to be the case.

The role of these house price changes in the MTM to private non-housing consumption is then assessed through the implementation of two alternative exercises in which the amplifying role of these assets is directly quantified. Results are very clear-cut and show that in the same countries where house prices are more sensitive to lending costs (Finland, Sweden and the UK) real house prices appear to have played an important role in the transmission of a monetary shock to household consumer spending during the sample period.

The above findings have important policy implications. In particular, they suggest that in European countries where the financial deregulation process was less intense, house prices might become a significant amplifying variable in the MTM and their role cannot be ignored in a proper assessment of the effects of interest rate changes on real activity. The recent upward competitive pressures witnessed in the mortgage markets in many continental countries and the subsequent increase in the housing lending to the household sector and mortgage products over the last few years are evident signs which indicate this. The role of these effects in these countries, for instance, could be captured with a more flexible empirical approach which allows for time-varying coefficients. The latter would not only reinforce the findings of this chapter, but would also allow for the identification of changes in the MTM during the period under examination.

APPENDIX 1 Data Description and Sources

Data are obtained from the International Financial Statistics (IFS) of the International Monetary Fund and the Business Sector Data Base (BSDB) of the OECD. The variables used in this chapter are the following:

GROSS DOMESTIC PRODUCT at constant prices. The quarterly real GDP series are taken from BSDB (series GDPV) and are seasonally adjusted.

CONSUMER PRICE INDEX. The series on the non-harmonised consumer price index (CPI) are obtained from IFS (series 64...ZF... for each national series) and the original base year is 1995=100.

PRIVATE FINAL CONSUMPTION at constant prices. The quarterly private consumption series are taken from BSDB (series CPV) and are seasonally adjusted. It includes goods and services, as well as durable products (such as cars, washing machines, and home computers), purchased by households. It excludes purchases of dwellings, but includes imputed rent for owner-occupied dwellings.

CREDIT TO PRIVATE DOMESTIC SECTOR. This series is taken from IFS (series 32D...ZF for each country).

COMMODITY PRICE INDEX. This is the world export commodity price index (U.S. Dollar) taken from IFS (series 000176AXDZF...).

NOMINAL EXCHANGE RATE. The nominal exchange rate refers to calculated units of domestic currency per D-Mark and are calculated by using the IFS data (series ...RH..ZF). For Ireland and Finland the relevant rate is Irish Punt/Sterling and Finnish Markka/US$, respectively.

SHORT-TERM INTEREST RATES. The nominal interest rate is the money market rate (IRS) and comes from IFS (series 60bZF...). Table A.1.1 shows the details of the precise definition of each series.

TABLE A.1.1 Short-Term Interest Rates Definition

	Definition
Belgium	Call Money Rate
Finland	Average Cost of CB Debt
Ireland	1 Month Fixed Rate
Italy	Money Market Rate
Netherlands	Call Money Rate
Spain	Call Money Rate
Sweden	Call Money Rate
United Kingdom	Overnight Inter-bank Rate

HOUSE PRICE INDEX. Table A.1.2 shows the sources of house price statistics together with their time span availability and description. The data refer to transaction prices recorded by private real estate associations and, in many cases, collected and published by central banks and central statistical offices in different countries. House price developments are at the national level and refer to sales prices of existing dwellings or, when not available, those of new dwellings. The indices do not directly take into account regional house price dynamics, which, as shown by several studies (see Meen and Andrew, 1998 for a review) can differ quite significantly. However, broad measures are used in order to represent the trends in average house prices in each country. Differences in the construction of these indices, in the characteristics of dwellings (i.e. mix of properties traded over time, quality changes, etc), in the composition of the dwelling stock within and across countries make quantitative attempts to draw strong international comparisons not very robust.

TABLE A.1.2 House Price Definitions and Sources

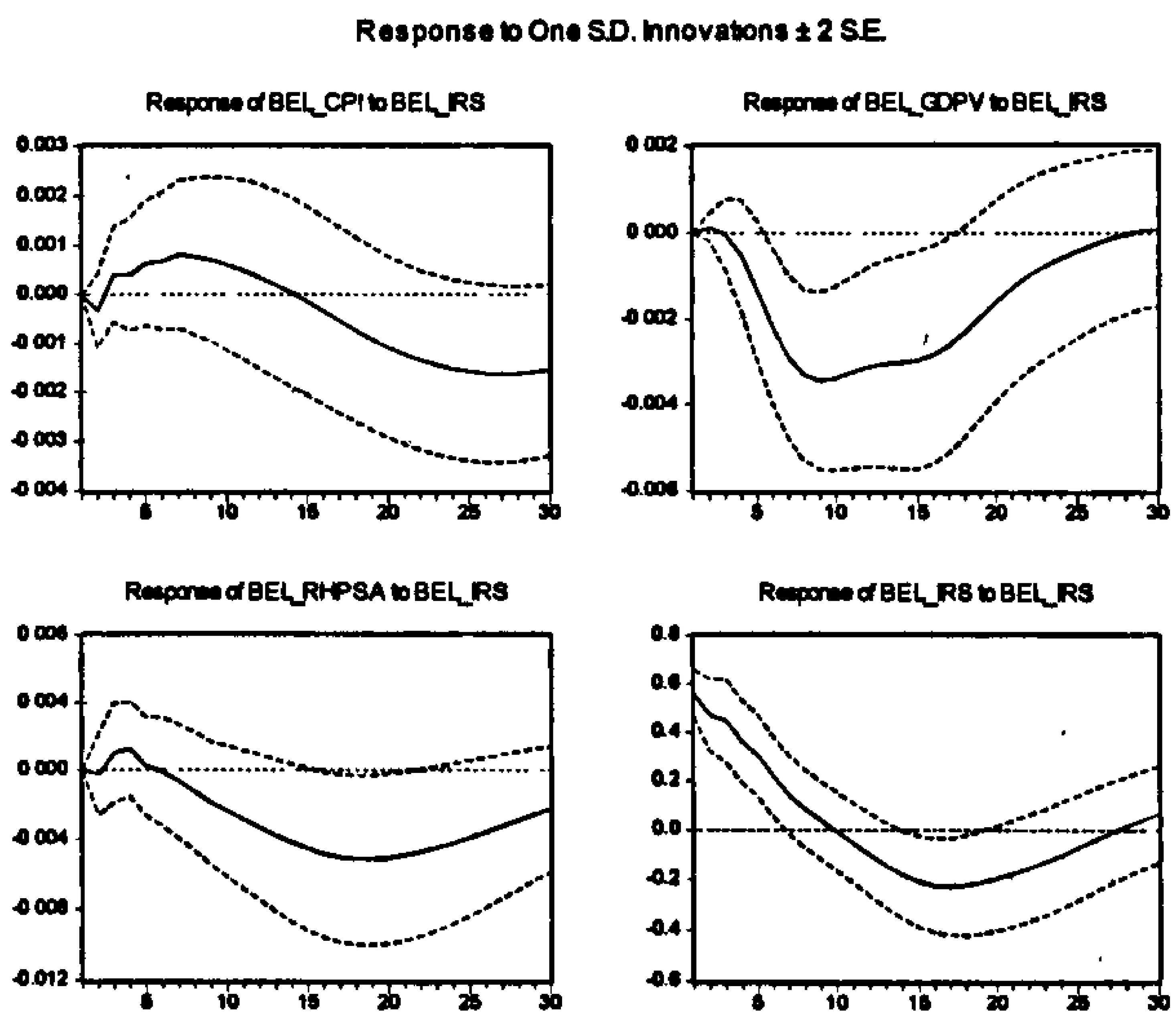
	Period	*Source*	*Definition/Notes*
Belgium	81Q1-99Q4	AN-Hyp Bruxelles, (Antwerpse Hypotheek Bank) Antwerpen	Price index for small and middle houses
Finland	70Q1-00Q2	Bank of Finland Huoneistokeskus	Average price for existing flats per square meter
Ireland	78Q1-00Q3	Department of the Environment and Local Government	Average gross prices of new (and second-hand) houses for which loans were approved by "All Agencies"
Italy	72S1-98S2	Iacoviello (2000) - "Il Consulente Immobiliare"	Residential property index
Netherlands	77Q1-99Q4	NVM (50-60% of total transactions)	Average Selling Price of existing single and multy-family houses
Spain	87Q1-00Q1	INE – Ministerio de Economia y Hacienda	Residential Property Price Index per Square Meter
Sweden	72Q1-99Q1	Iacoviello (2000) – Central Statistical Office (CSO)	House Price Index as weighted mean of primary and leisure homes
United Kingdom	74Q2-00Q1	Department of the Environment, Transport and the Regions (DETR)	UK mix-adjusted house price index, all dwellings; based on survey of all lenders

The semi-annual original series for Italy was transformed into quarterly series by using the RATS interpolation procedure INTERPOL.SRC, which allows for the estimation of different ARIMA models. A number of alternative models have been estimated which all show that different interpolation approaches were not affecting the main results of the VAR analysis. The results shown in the main text are based on an ARIMA (1,1,0). All series have been deflated using the CPI index. For the United Kingdom the RPIX (which does not include mortgage interest rate payments) has been used.

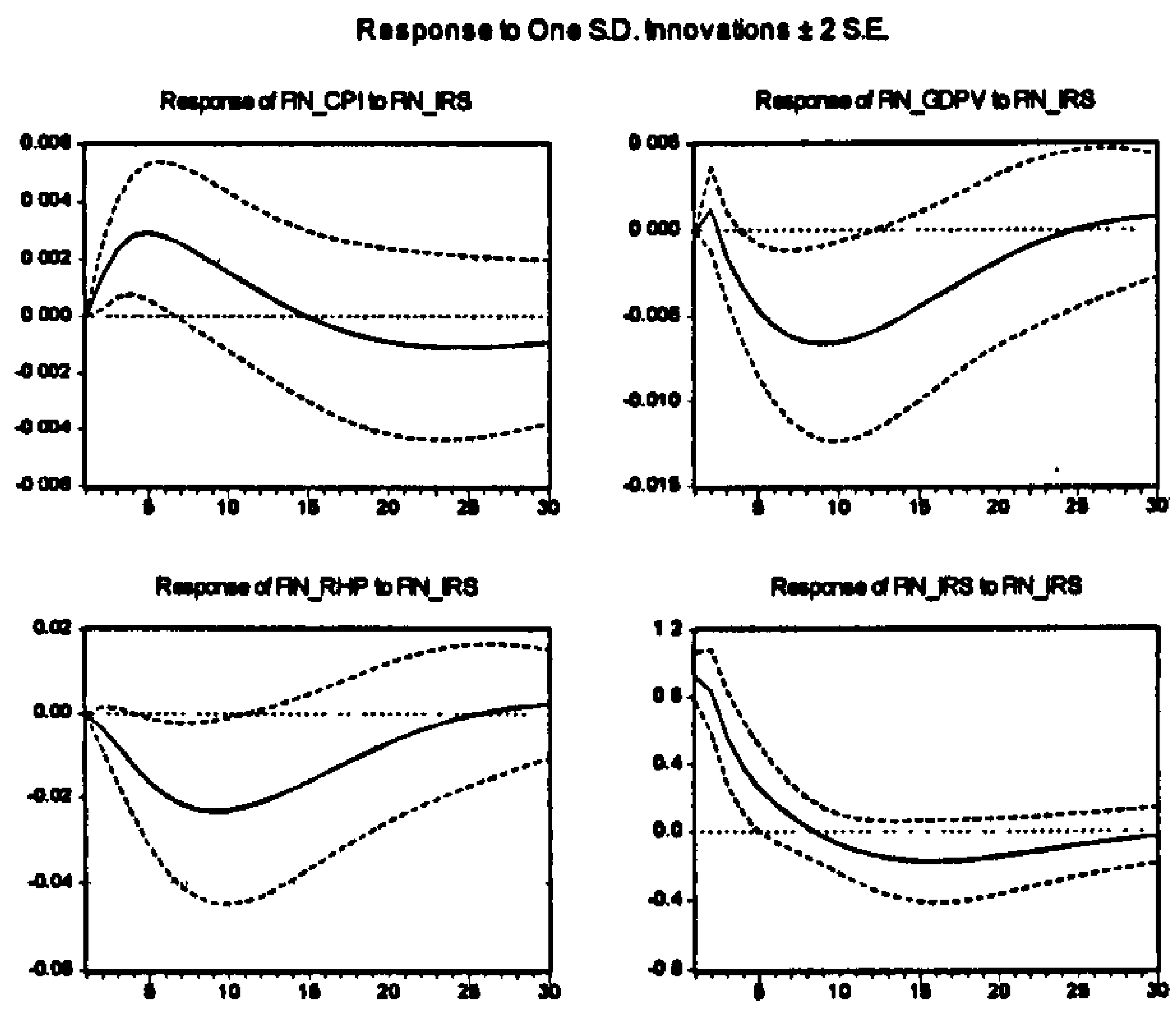
APPENDIX 2 *Estimation Results*

FIGURE A.2.1 **Impulse Responses to Monetary Shock in the Recursive Four-Variable VAR Model** (dotted lines: two standard error bands)

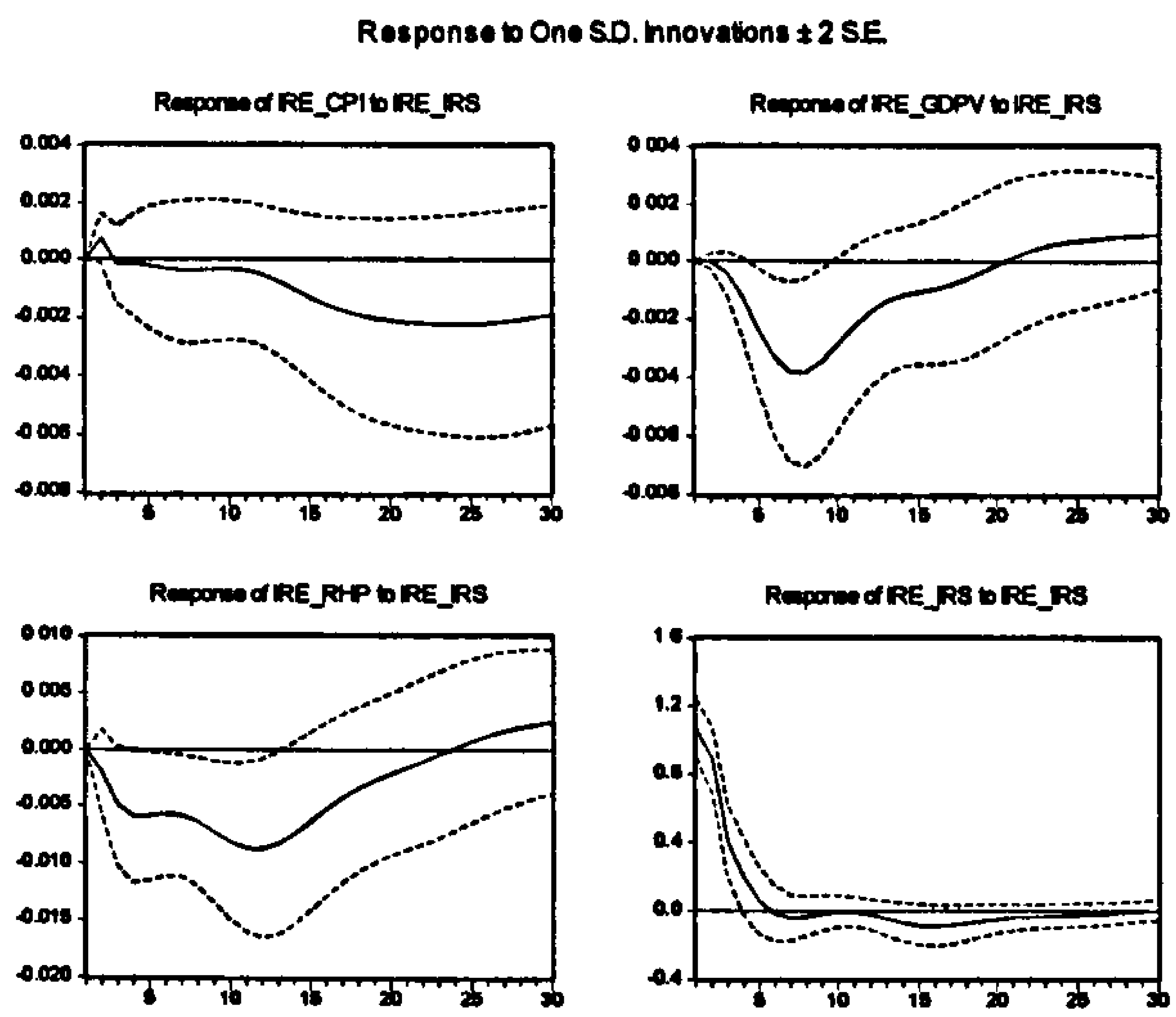
Belgium



Finland



Ireland



United Kingdom

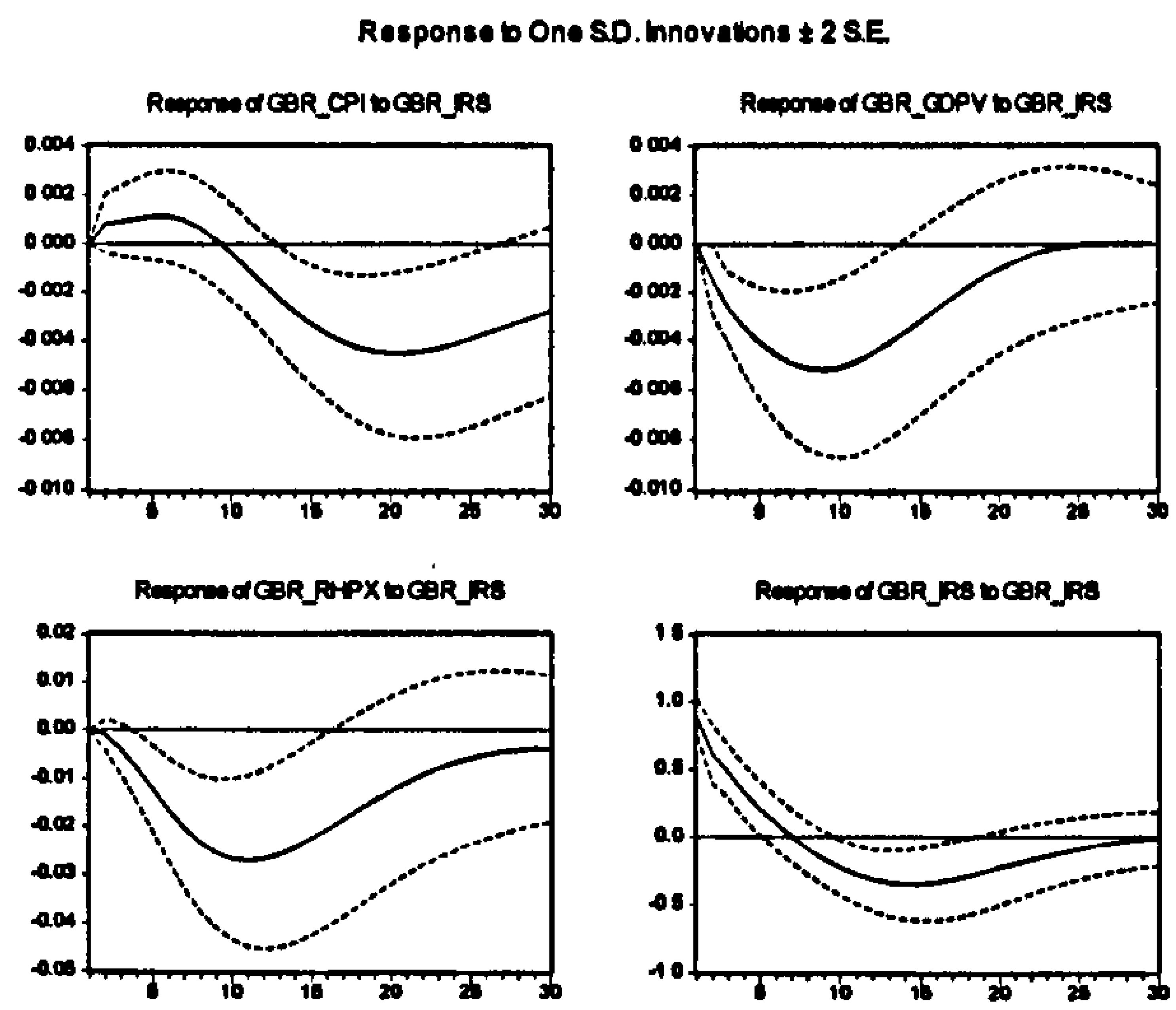
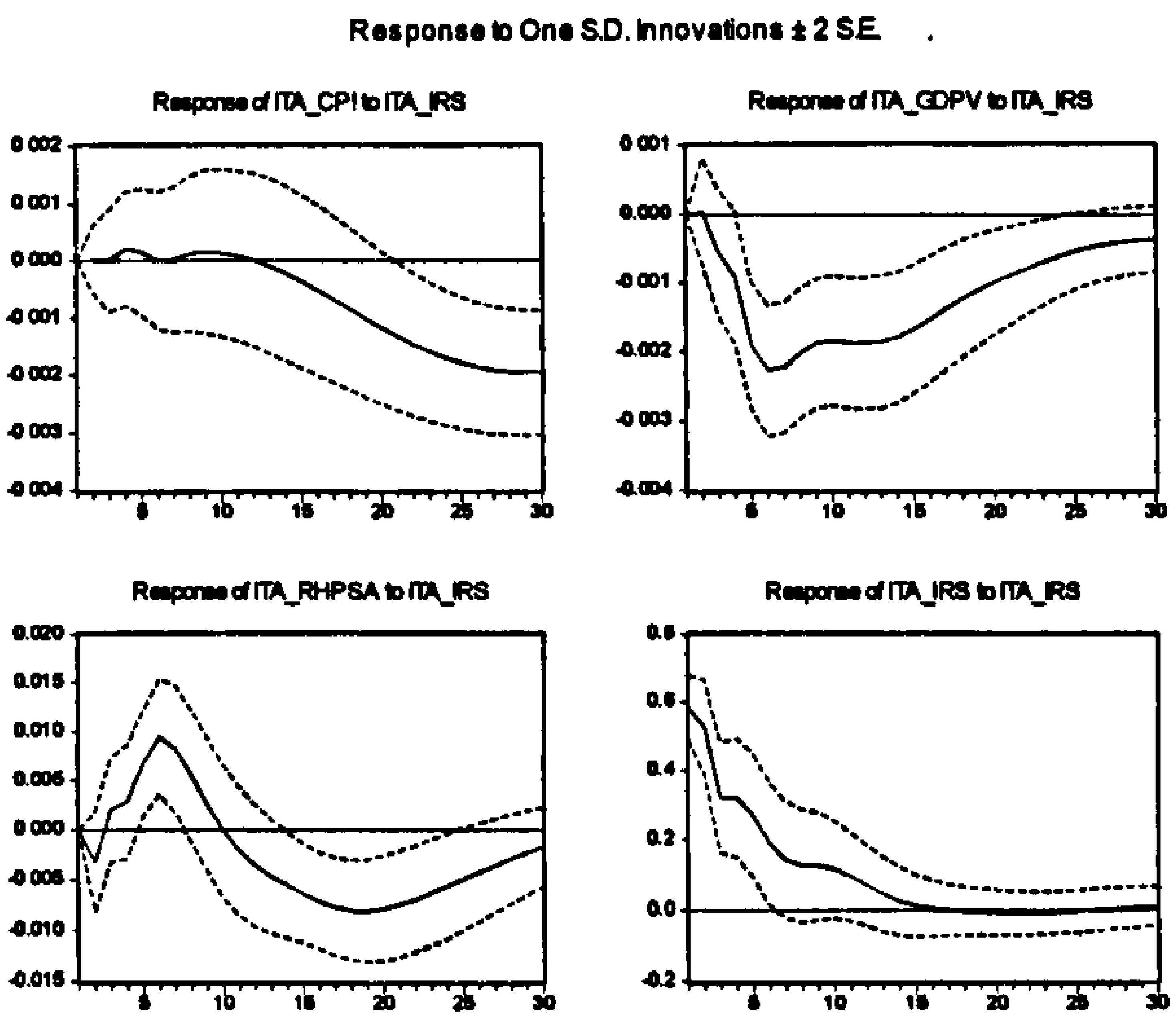
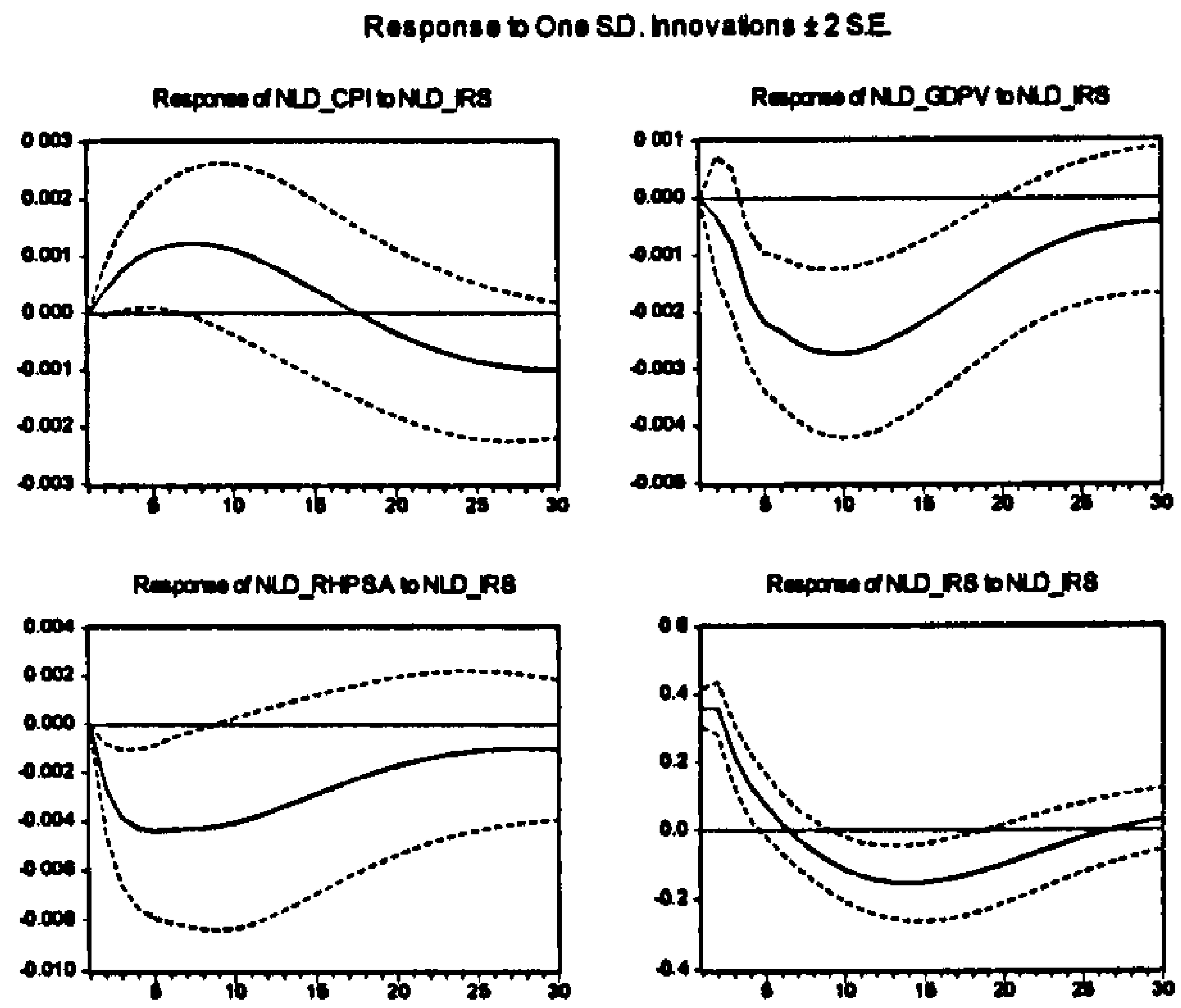


FIGURE A.2.1 **Impulse Responses to Monetary Shock in the Recursive Four-Variable VAR Model** (dotted lines: two standard error bands)

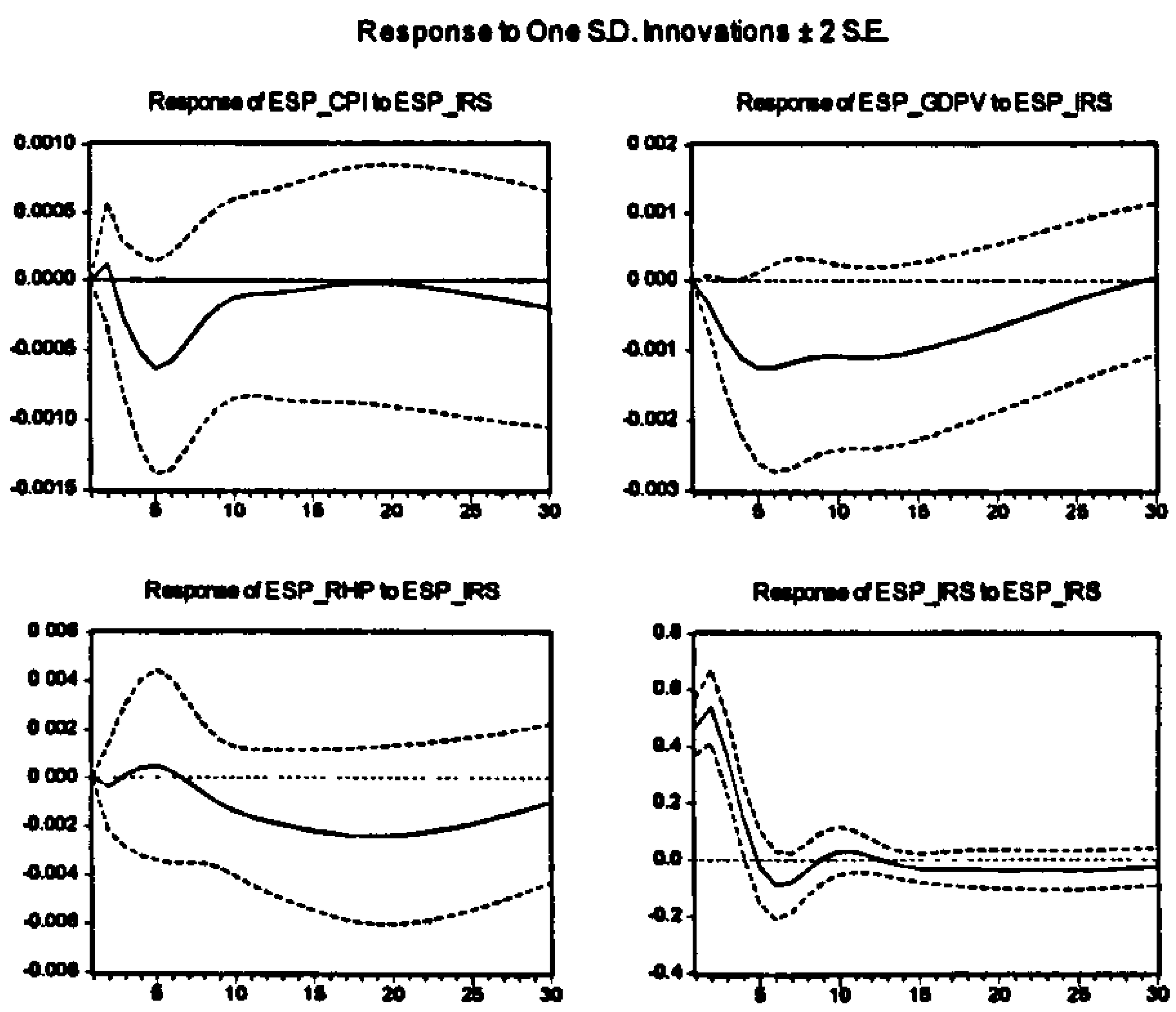
Italy



Netherlands



Spain



Sweden

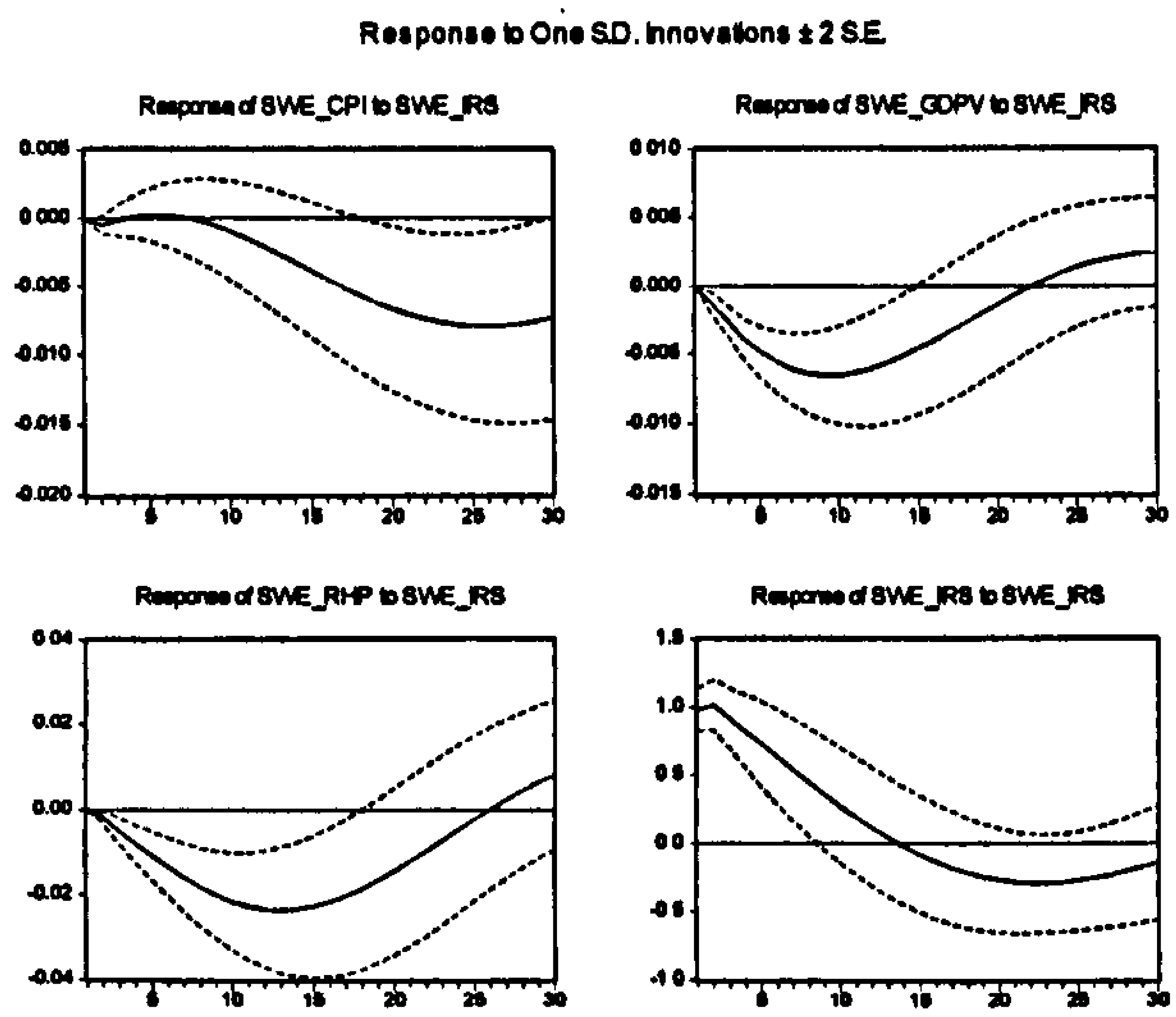
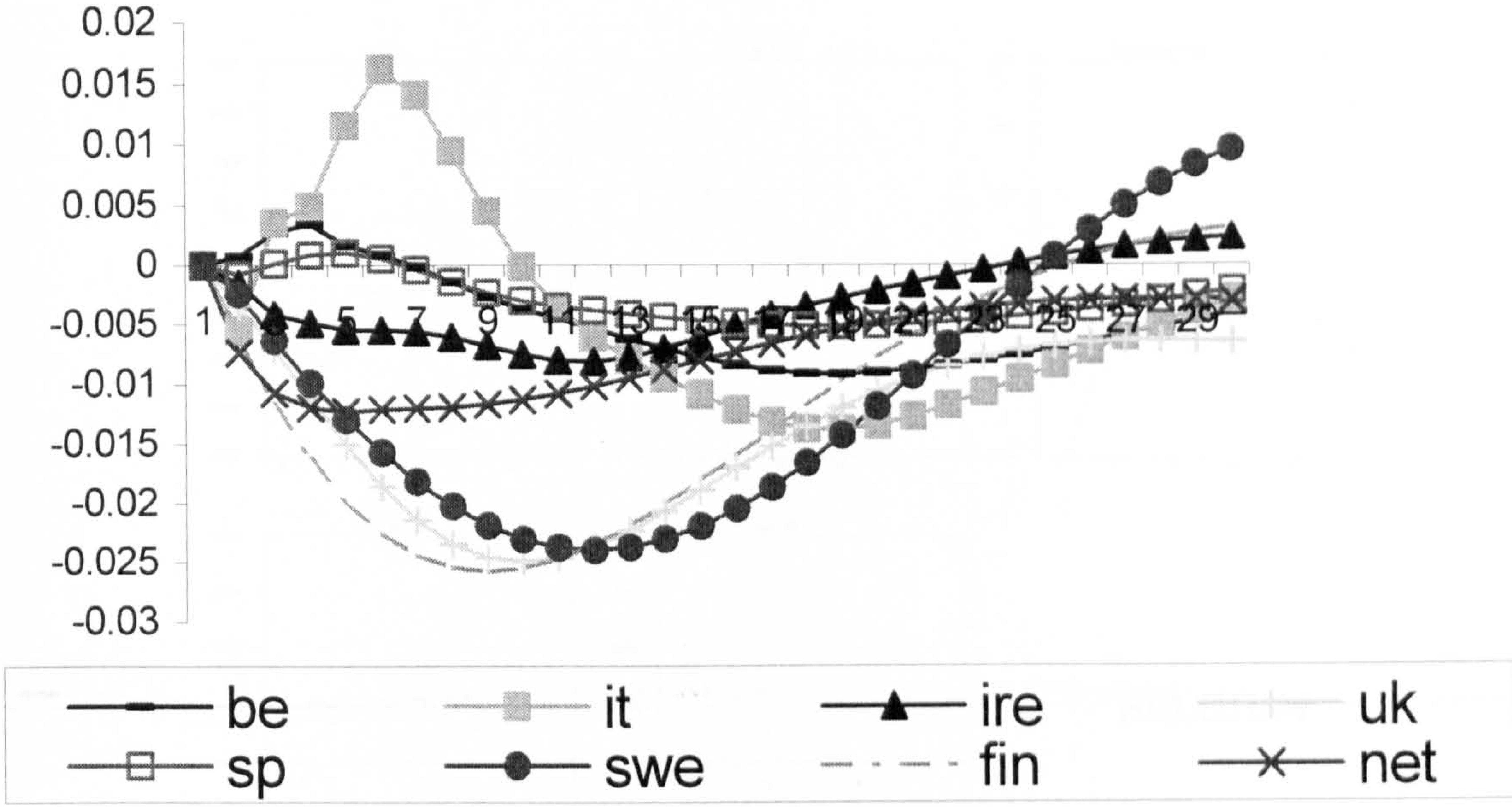
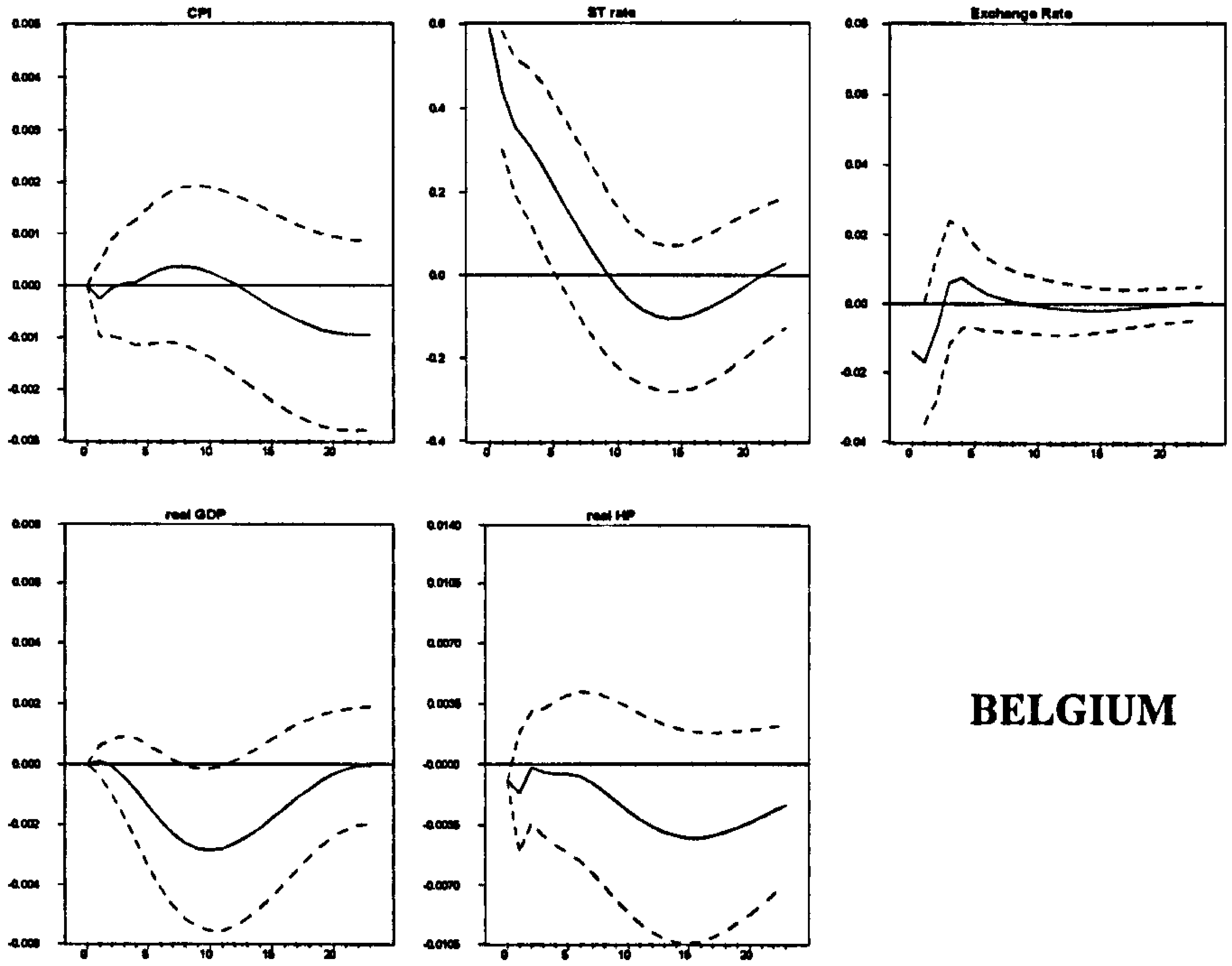


FIGURE A.2.1a **Real House Price Response to a Normalised Monetary Shock in the Recursive Four-Variable VAR Model**

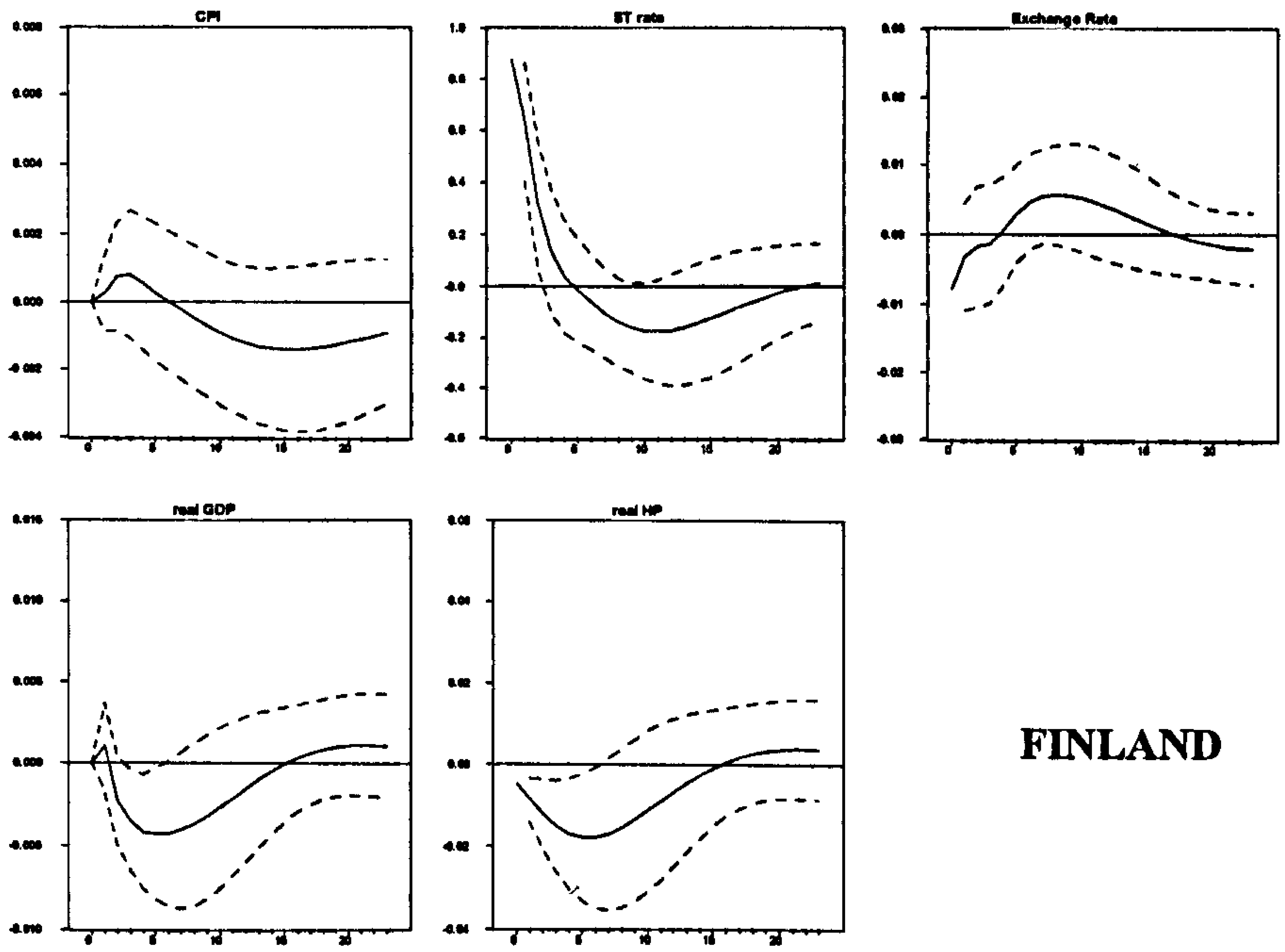


Notes: Impulse response functions of the real house price to a temporary normalized (100-point) short-term interest rate shock estimated in the recursive four-variable system. Be = Belgium, It = Italy, Ire = Ireland, UK = United Kingdom, Sp = Spain, Swe = Sweden, Fin = Finland and Net = the Netherlands.

FIGURE A.2.2 **Impulse Responses to Monetary Policy in the Five-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)

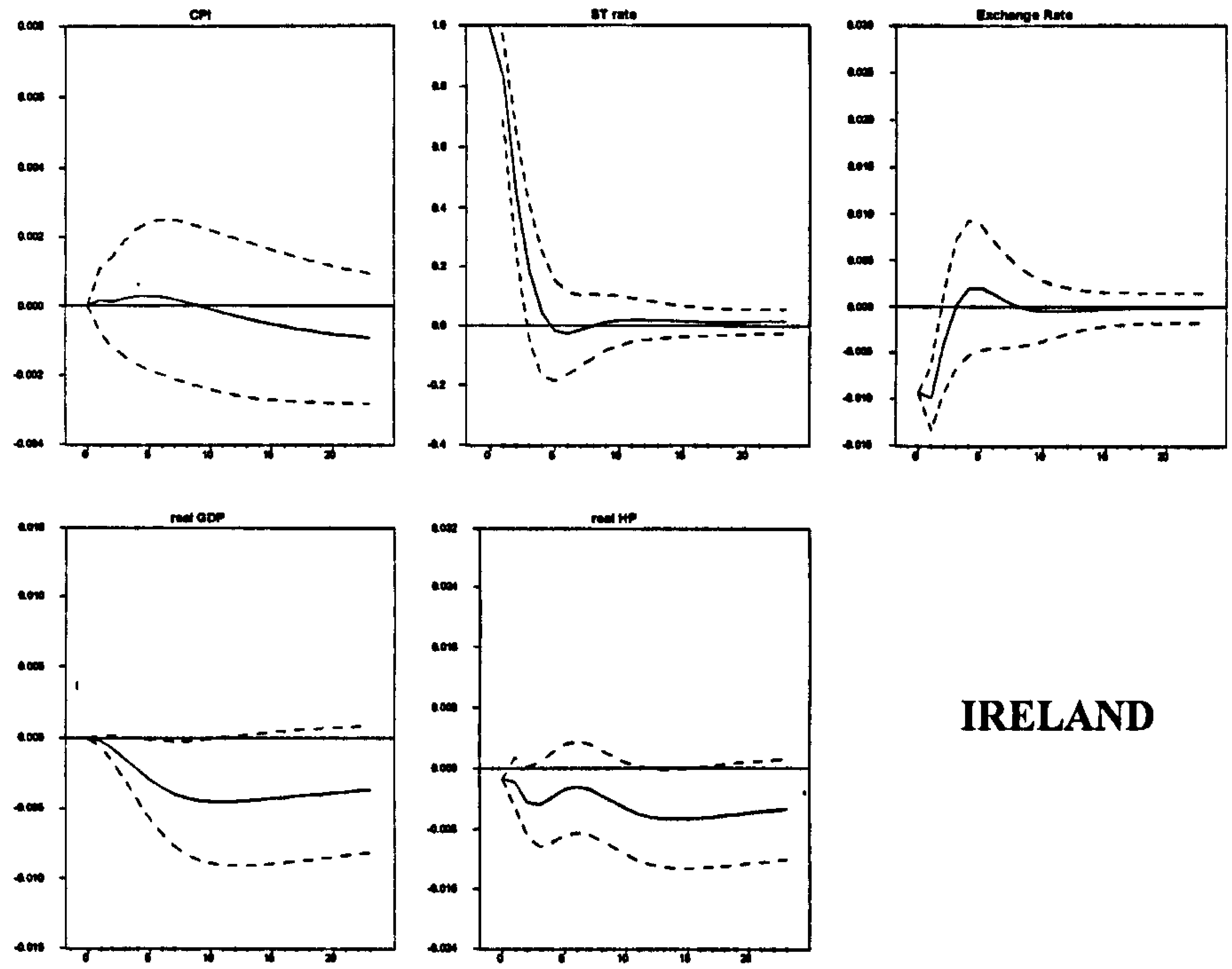


BELGIUM

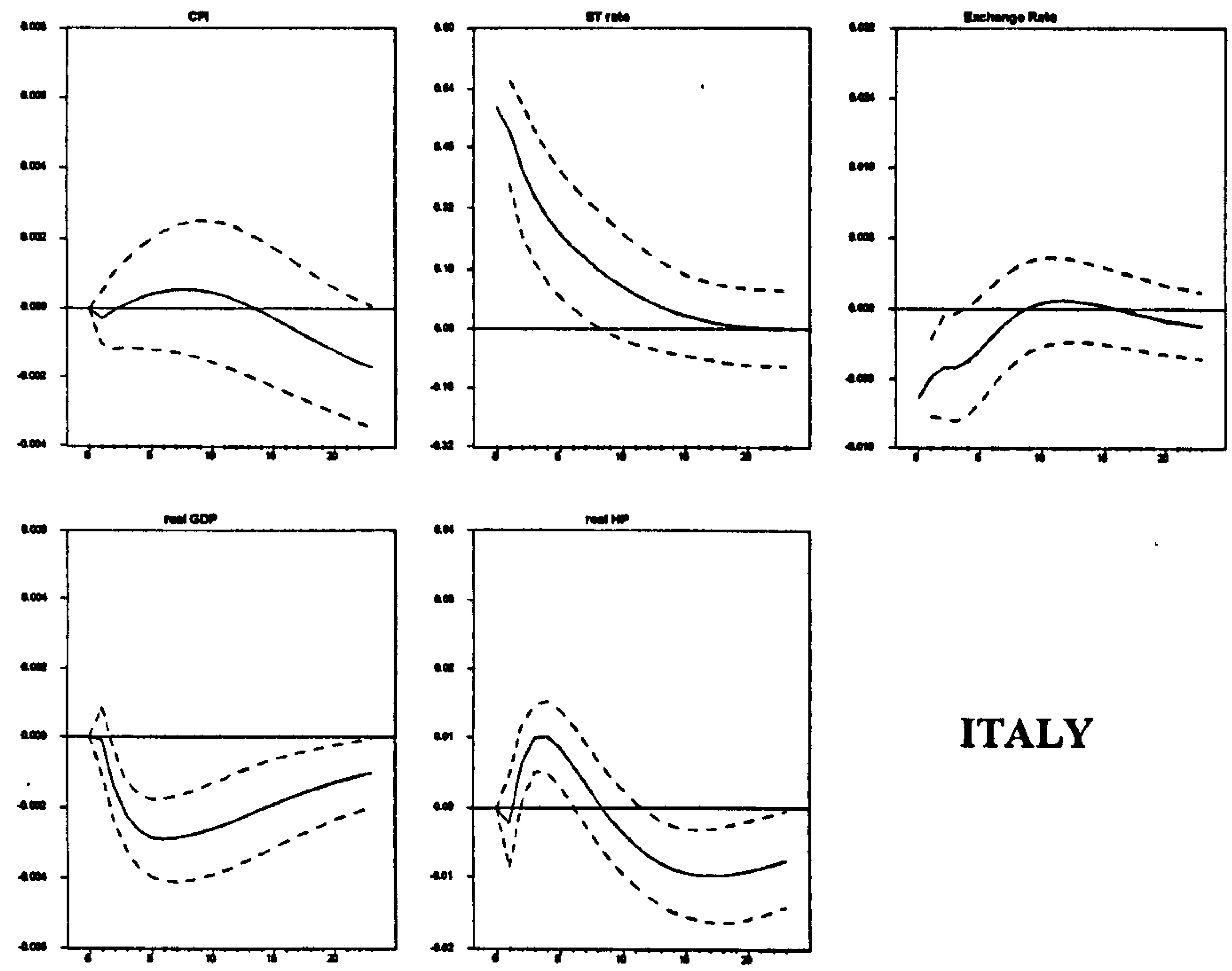


FINLAND

FIGURE A.2.2 **Impulse Responses to Monetary Shock in the Five-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)

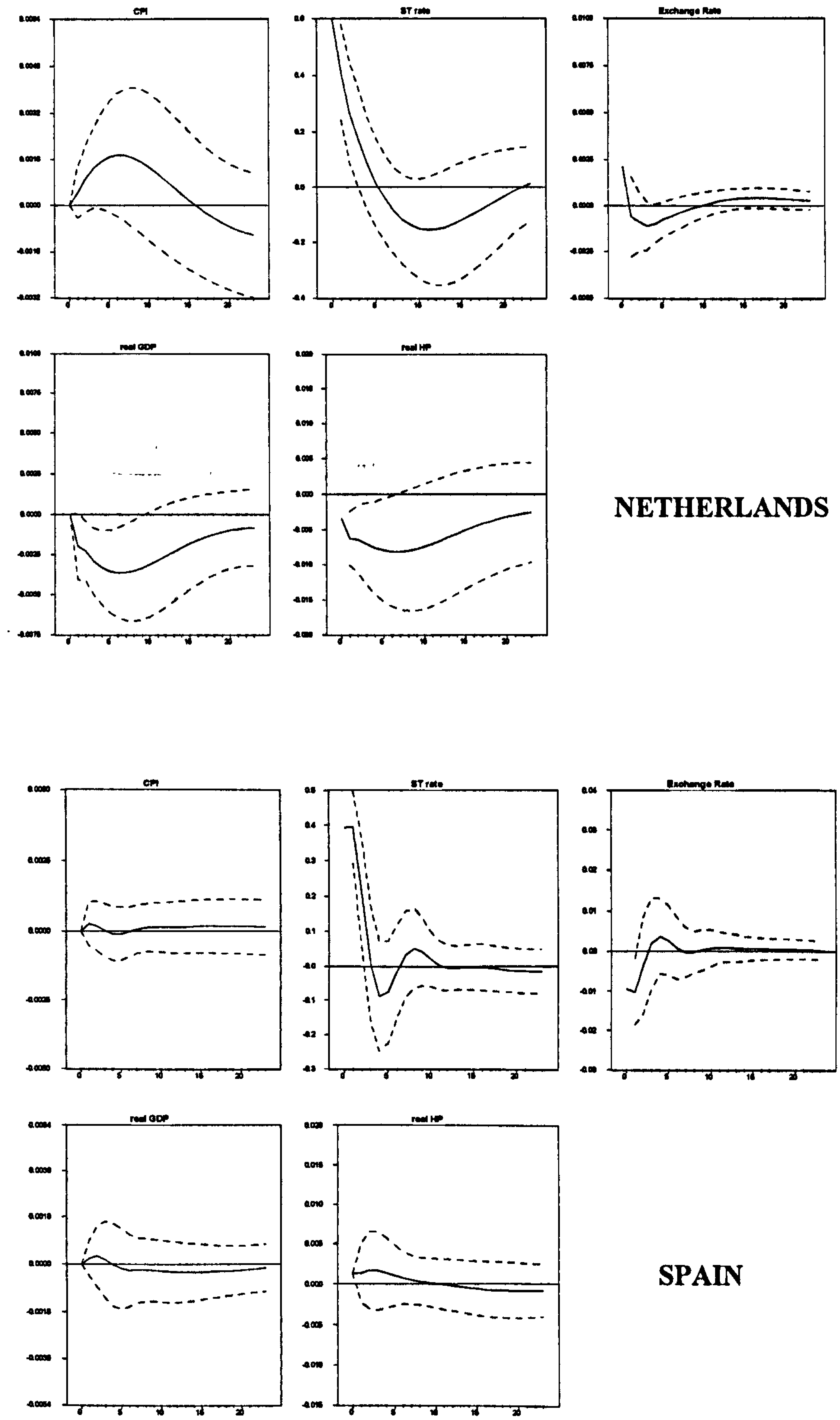


IRELAND



ITALY

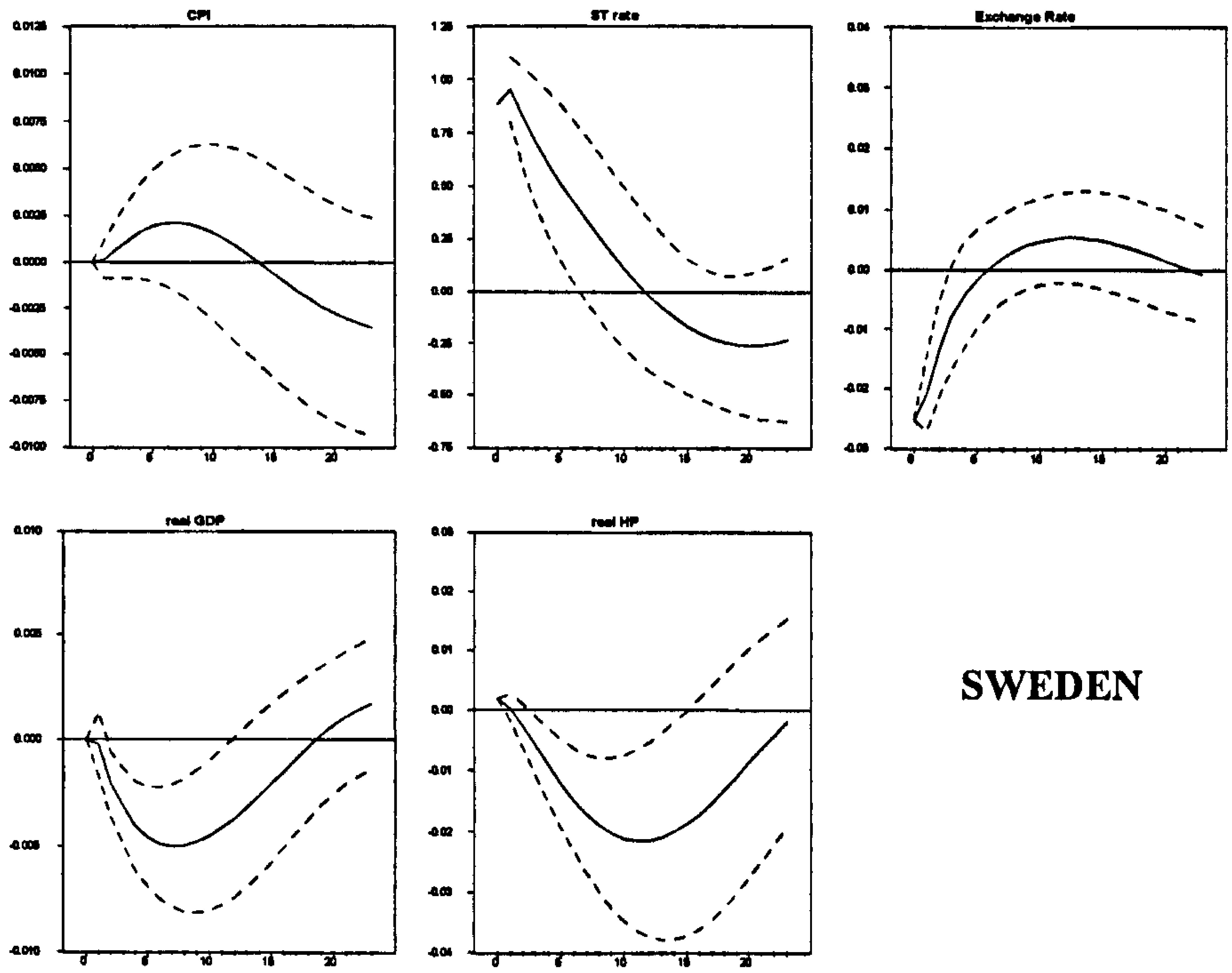
FIGURE A.2.2 **Impulse Responses to Monetary Shock in the Five-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)



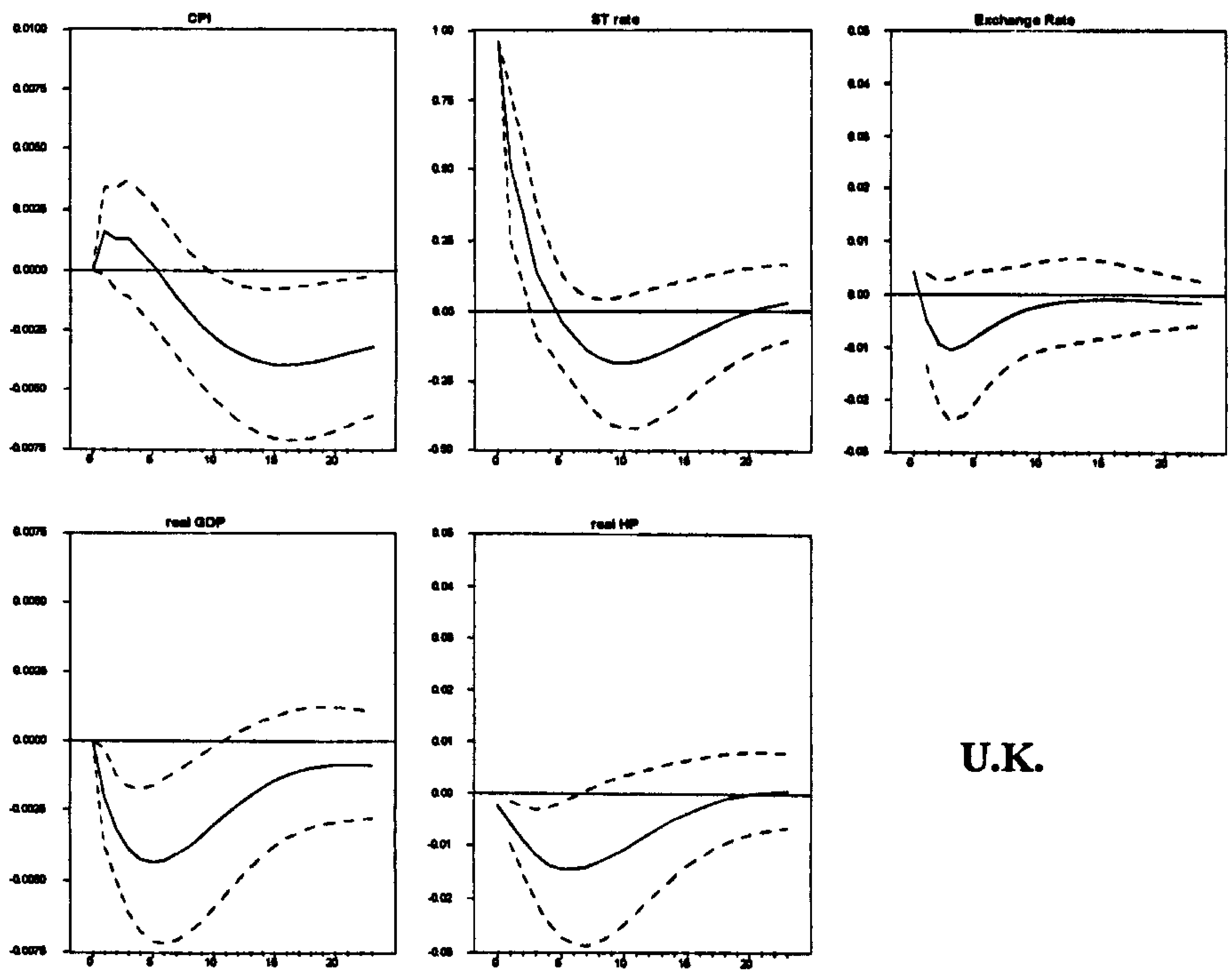
NETHERLANDS

SPAIN

FIGURE A.2.2 **Impulse Responses to Monetary Shock in the Five-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)



SWEDEN



U.K.

FIGURE A.2.3 **Forecast Error Variance Decomposition for Real House Price – Five-Variable SVAR**

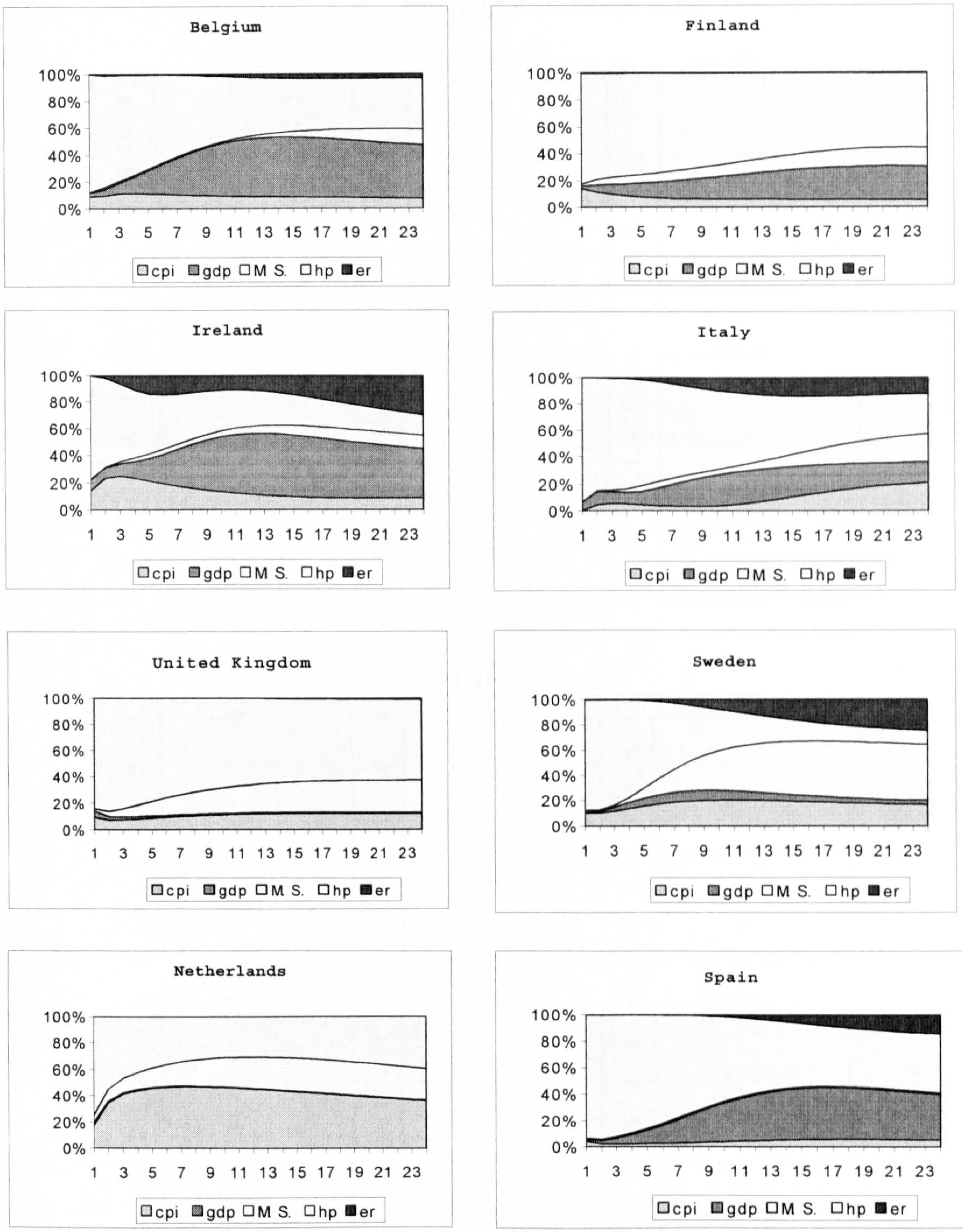
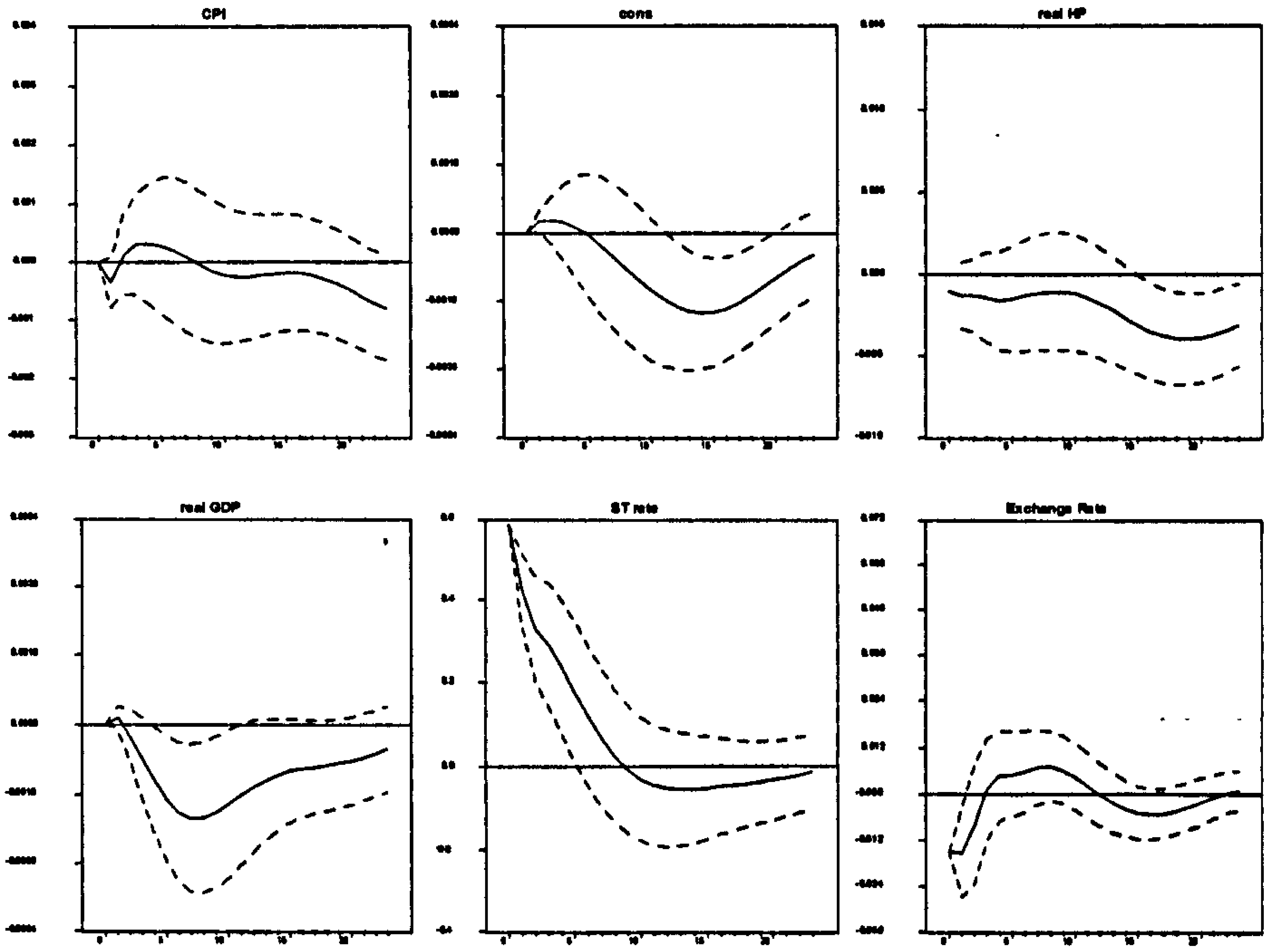


FIGURE A.2.4 **Impulse Responses to Monetary Shock in the Six-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)

BELGIUM



FINLAND

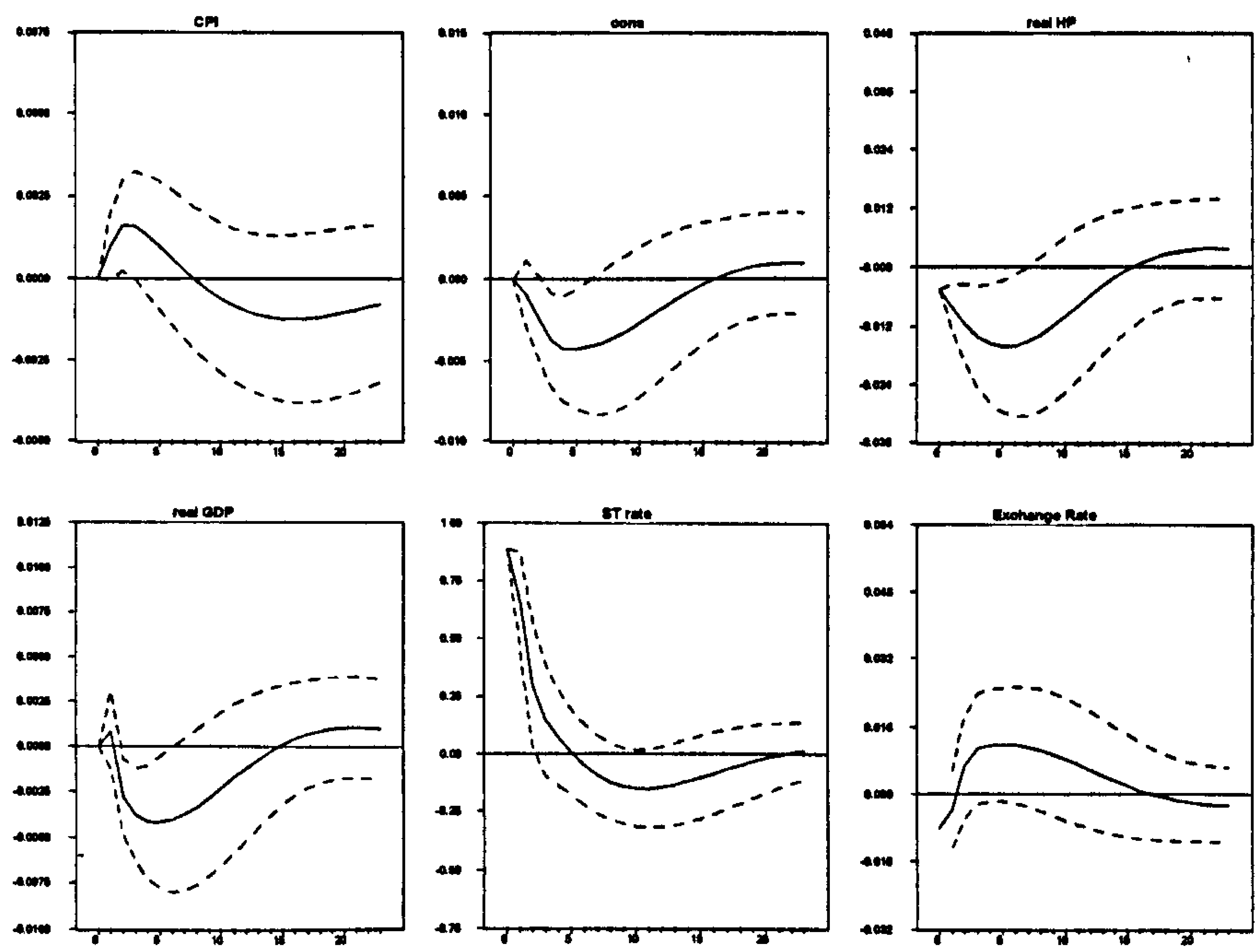
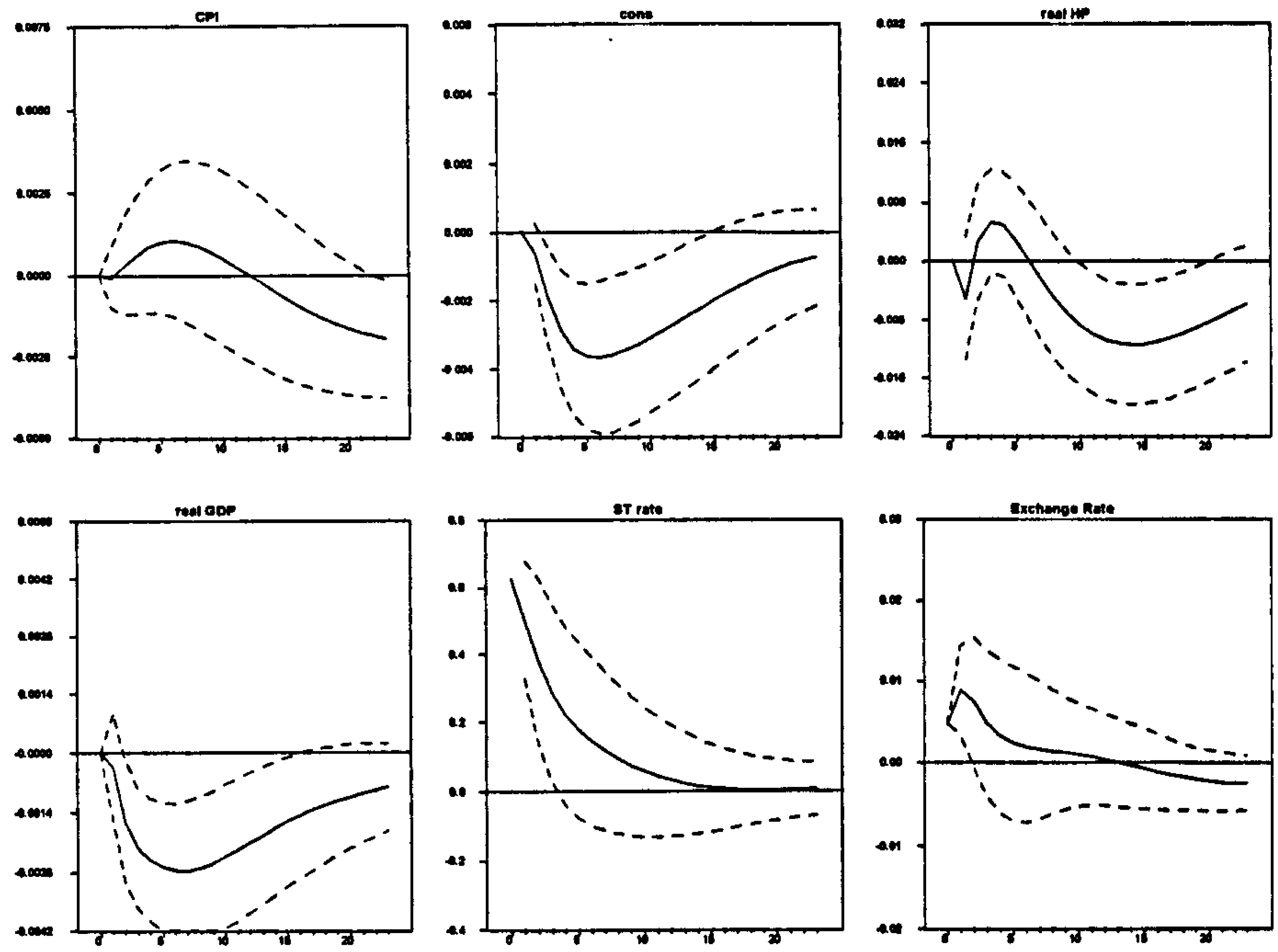


FIGURE A.2.4 **Impulse Responses to Monetary Shock in the Six-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)

ITALY



NETHERLANDS

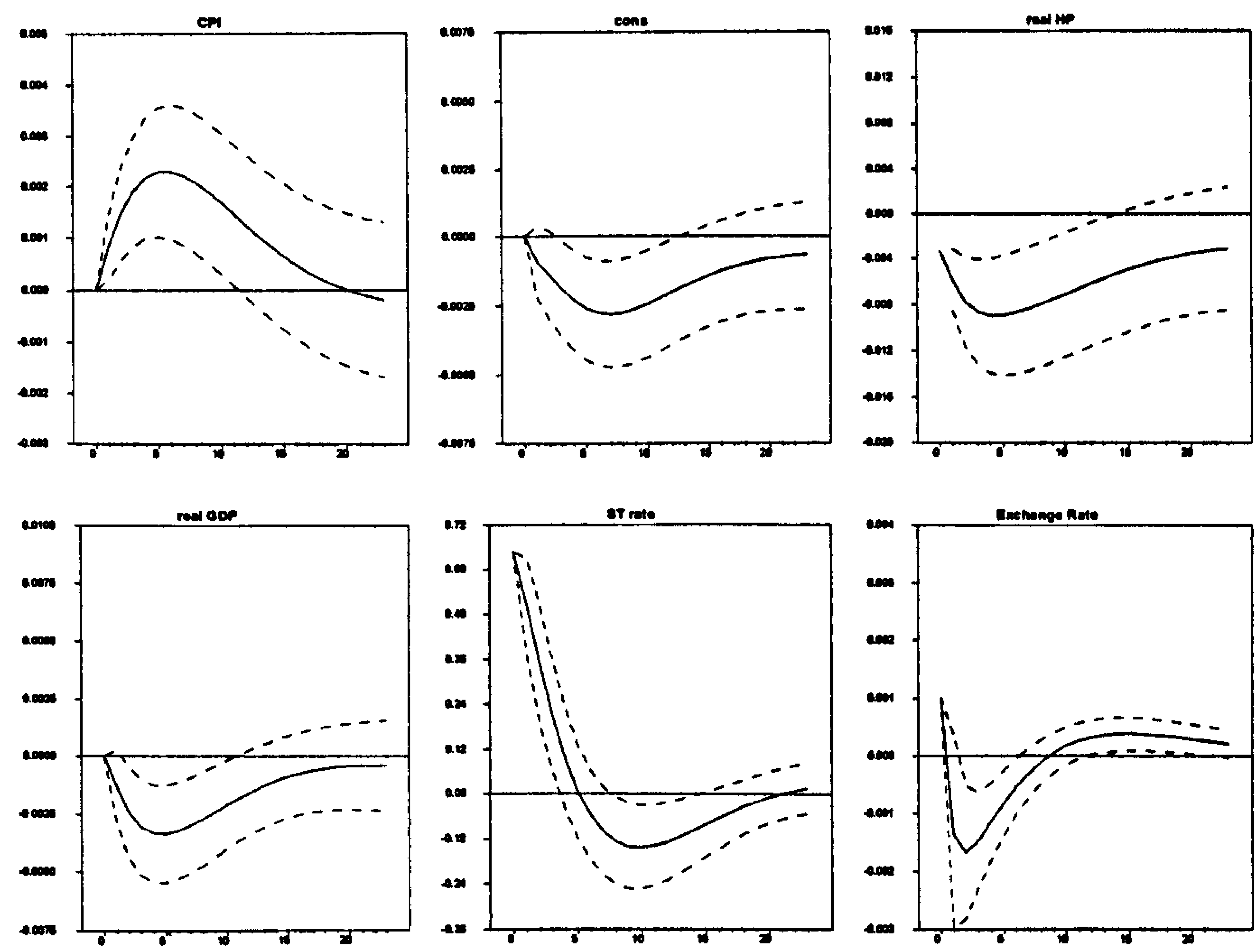
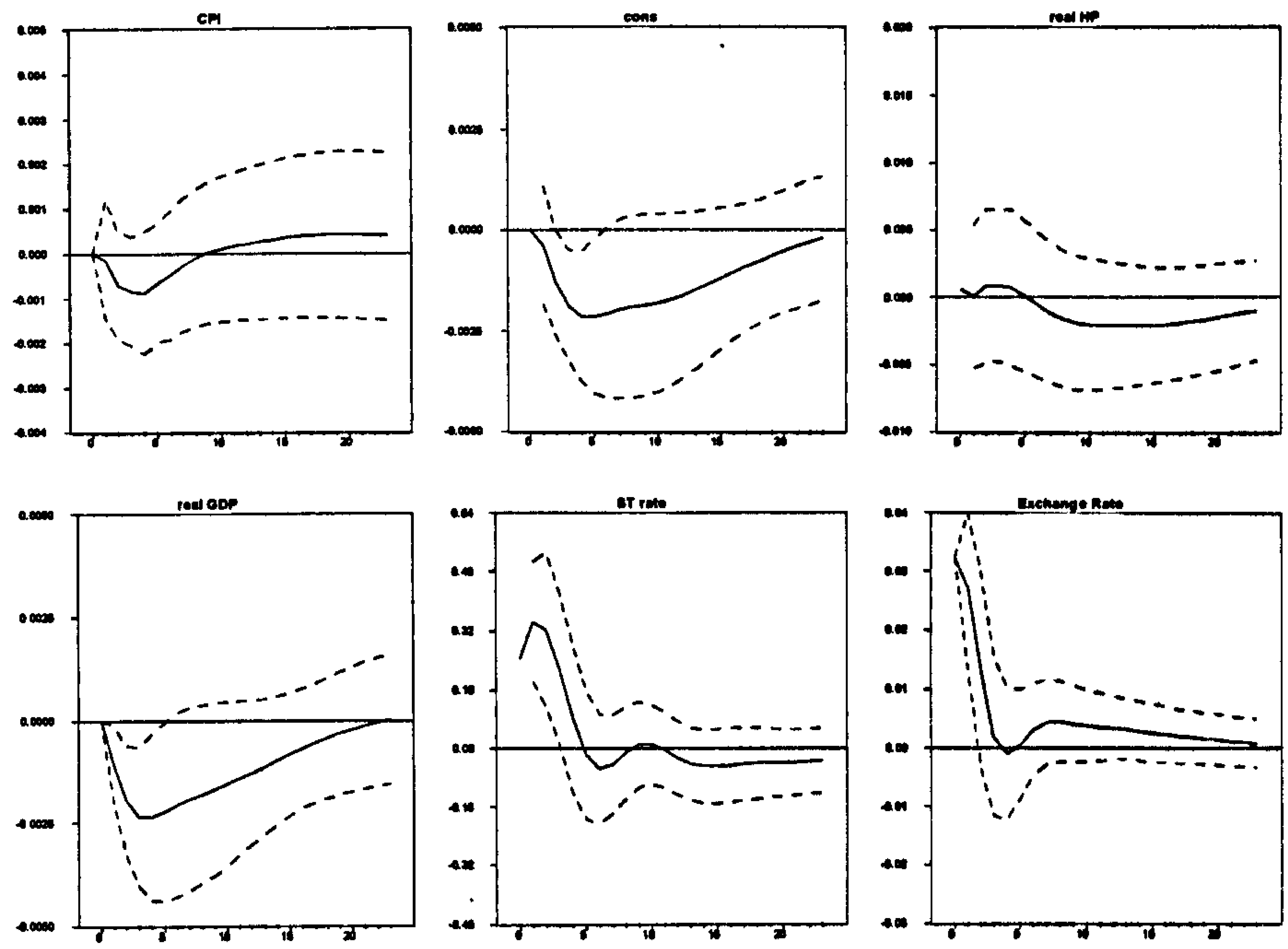


FIGURE A.2.4 **Impulse Responses to Monetary Shock in the Six-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)

SPAIN



SWEDEN

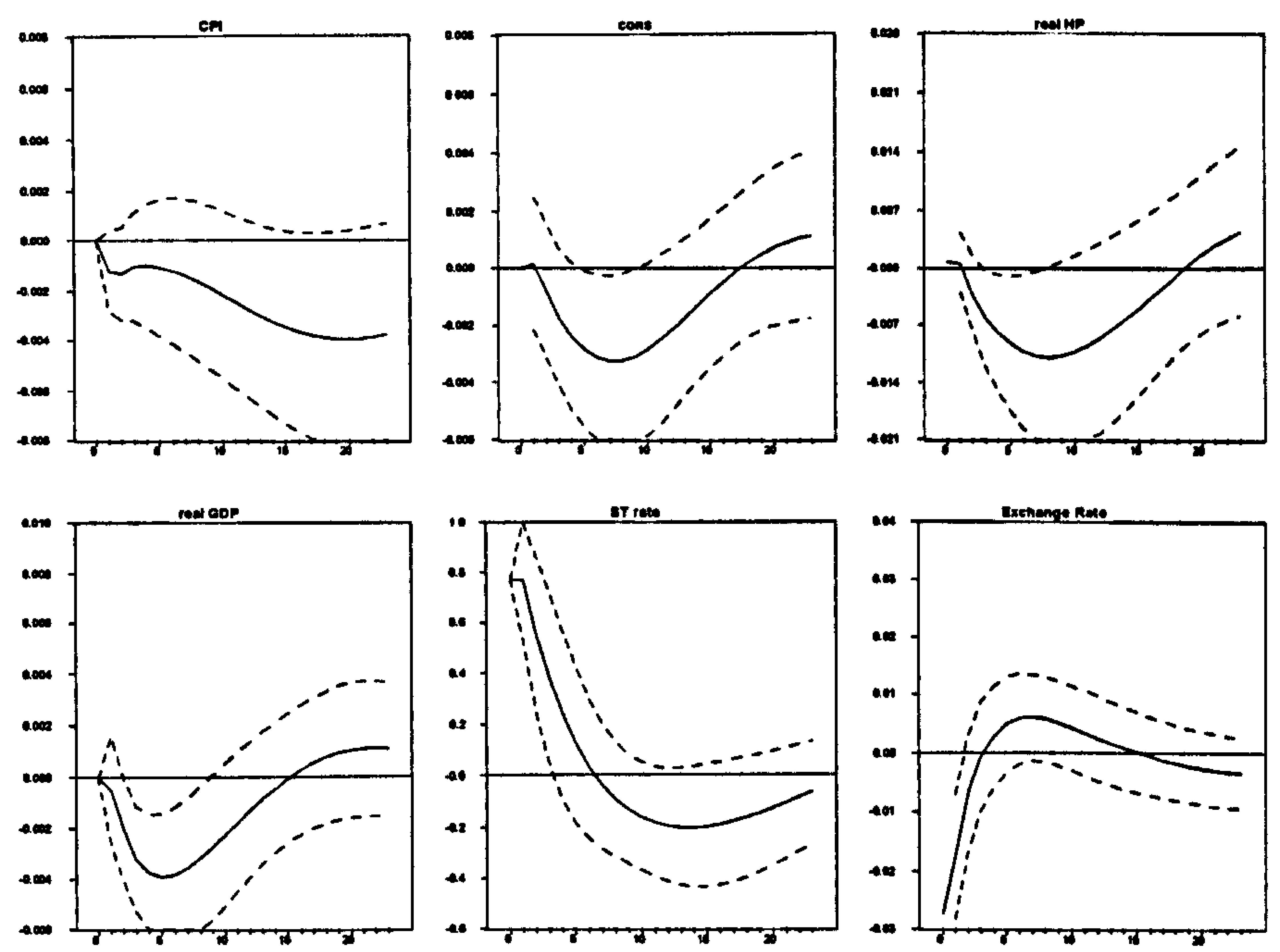


FIGURE A.2.4 **Impulse Responses to Monetary Shock in the Six-Variable SVAR** (dotted lines are two-standard deviation confidence intervals)

UNITED KINGDOM

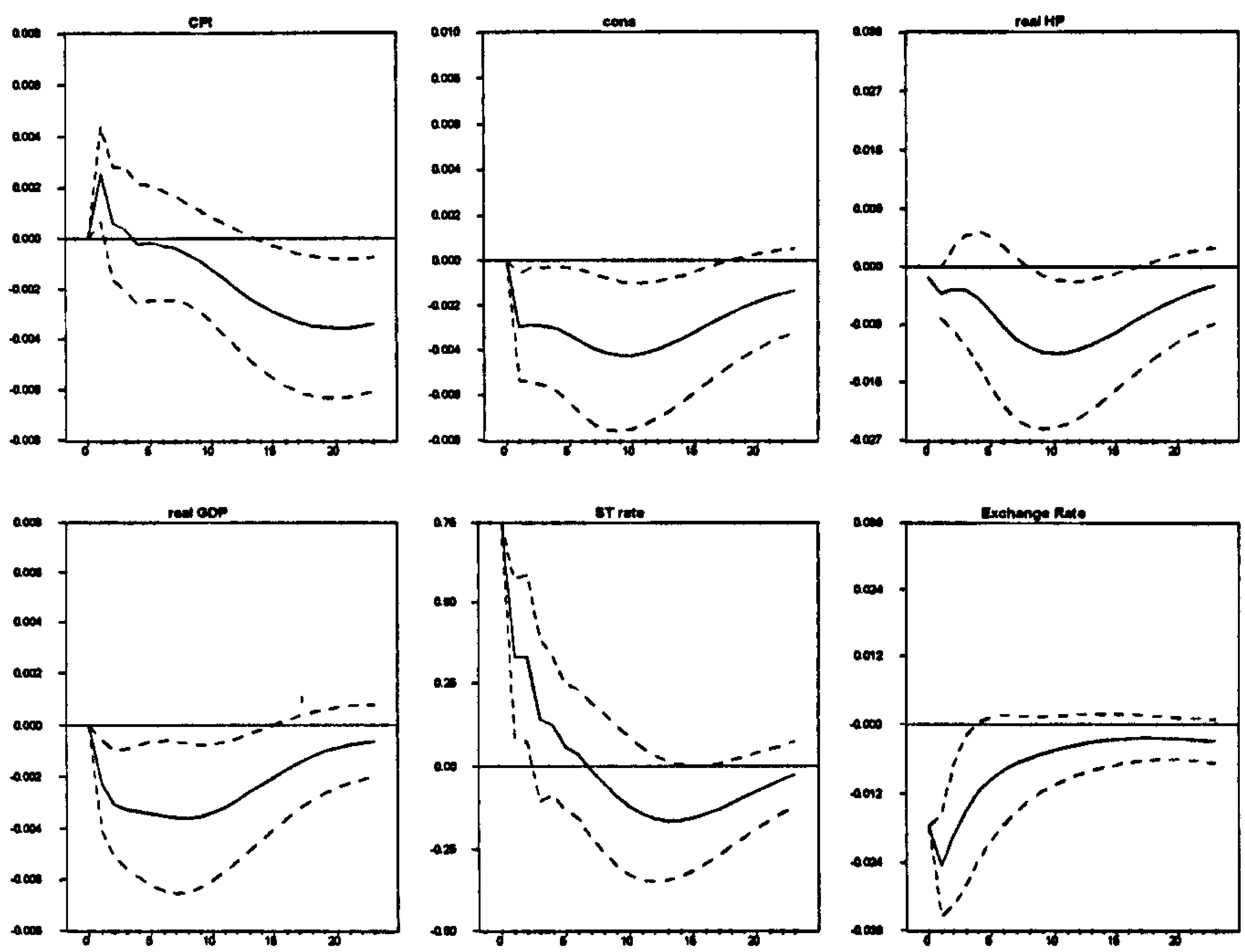
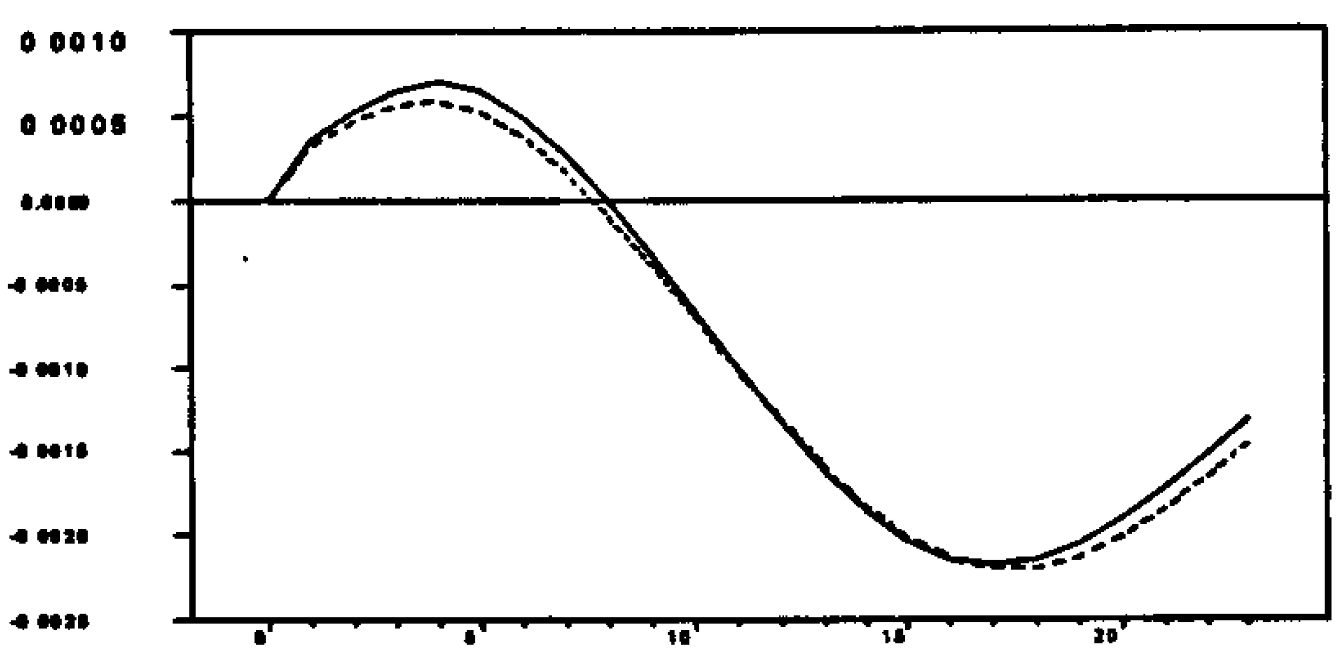
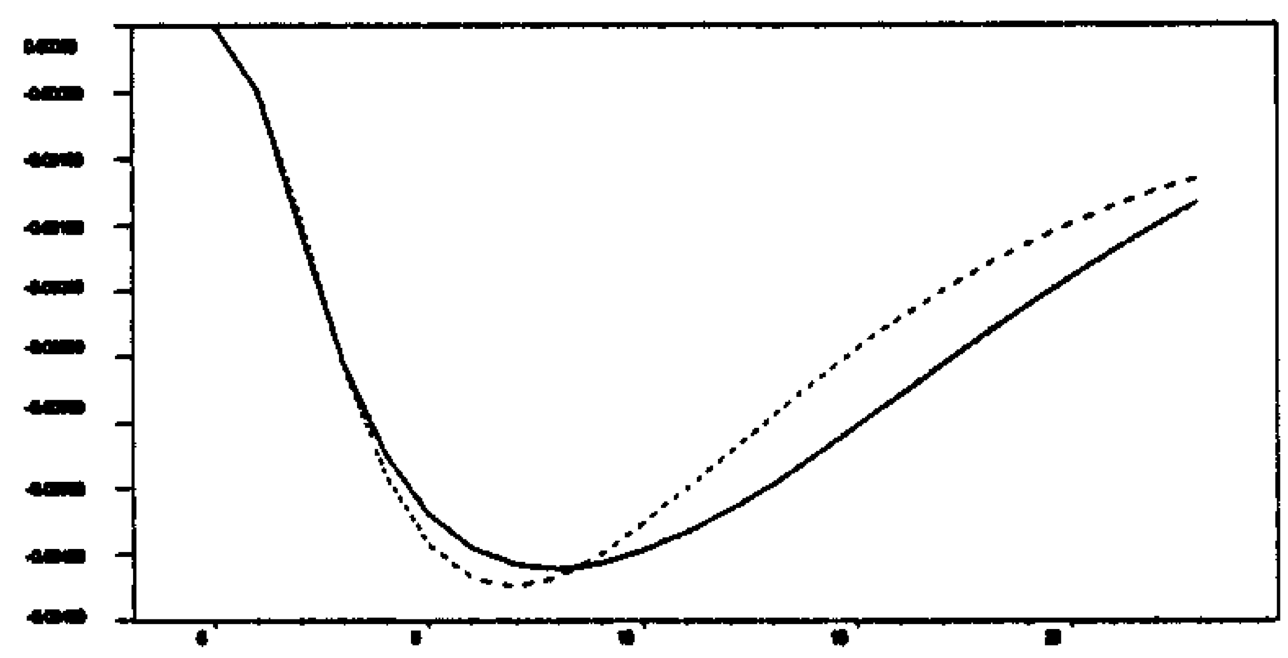


FIGURE A.2.5 **Response of Real Private Consumption to Monetary Shock - Counterfactual Simulation**
(solid line: baseline - dotted line: counterfactual scenario)

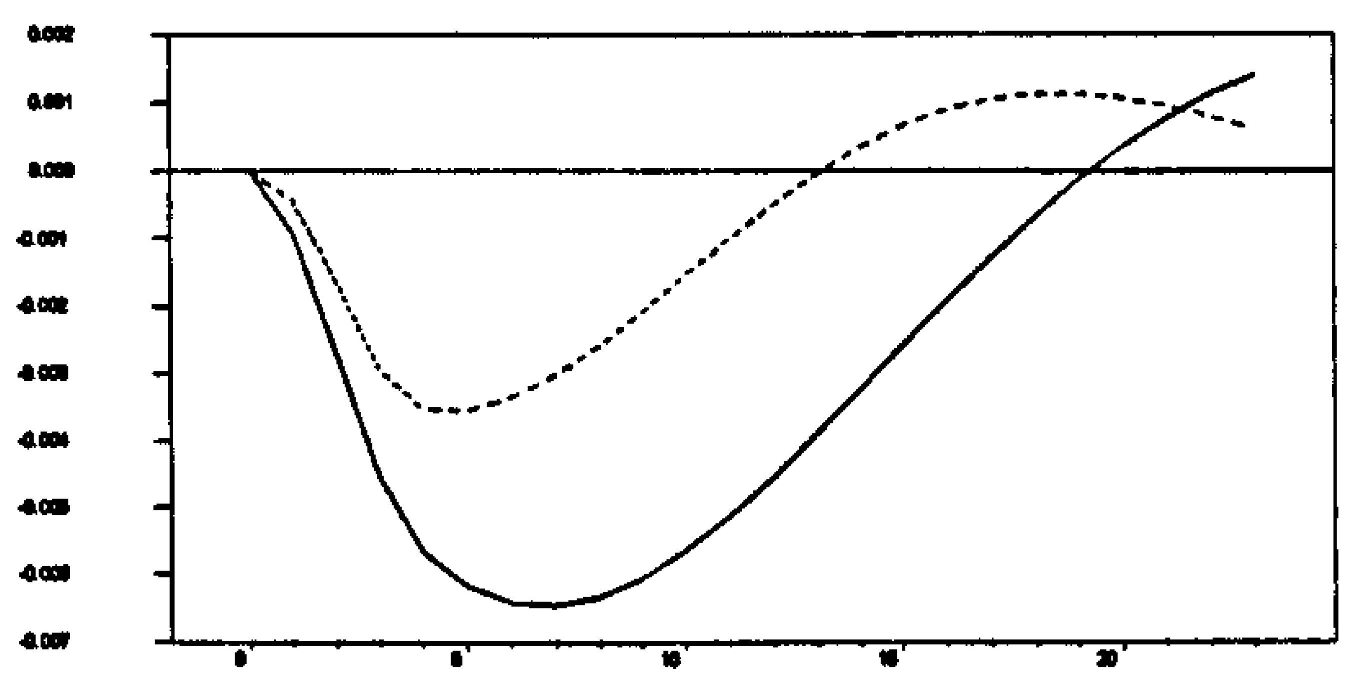
BELGIUM



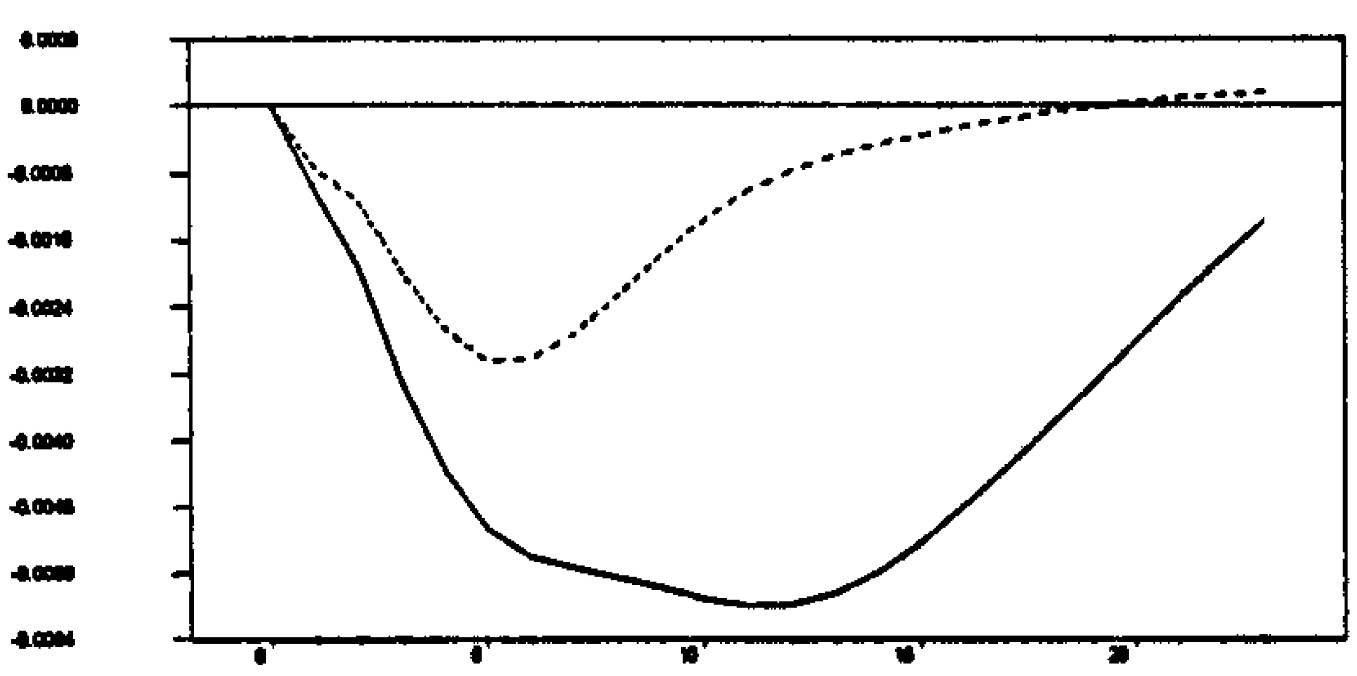
ITALY



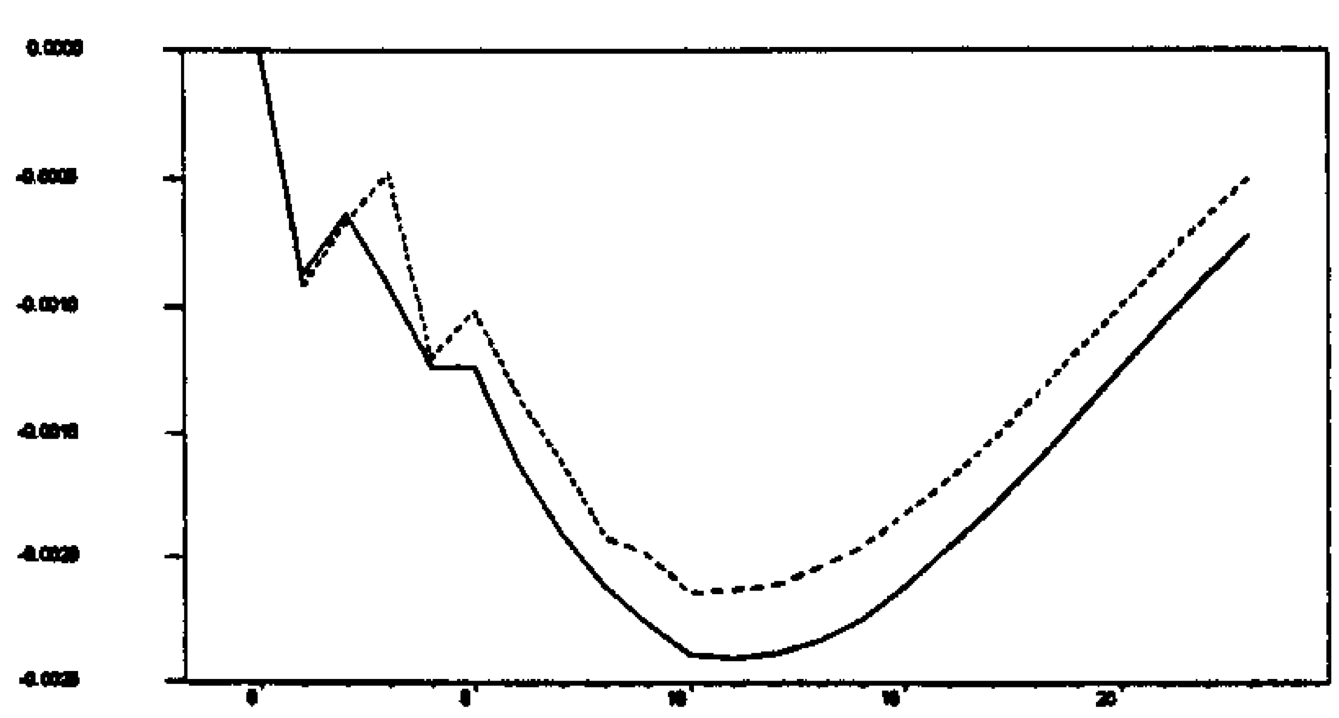
FINLAND



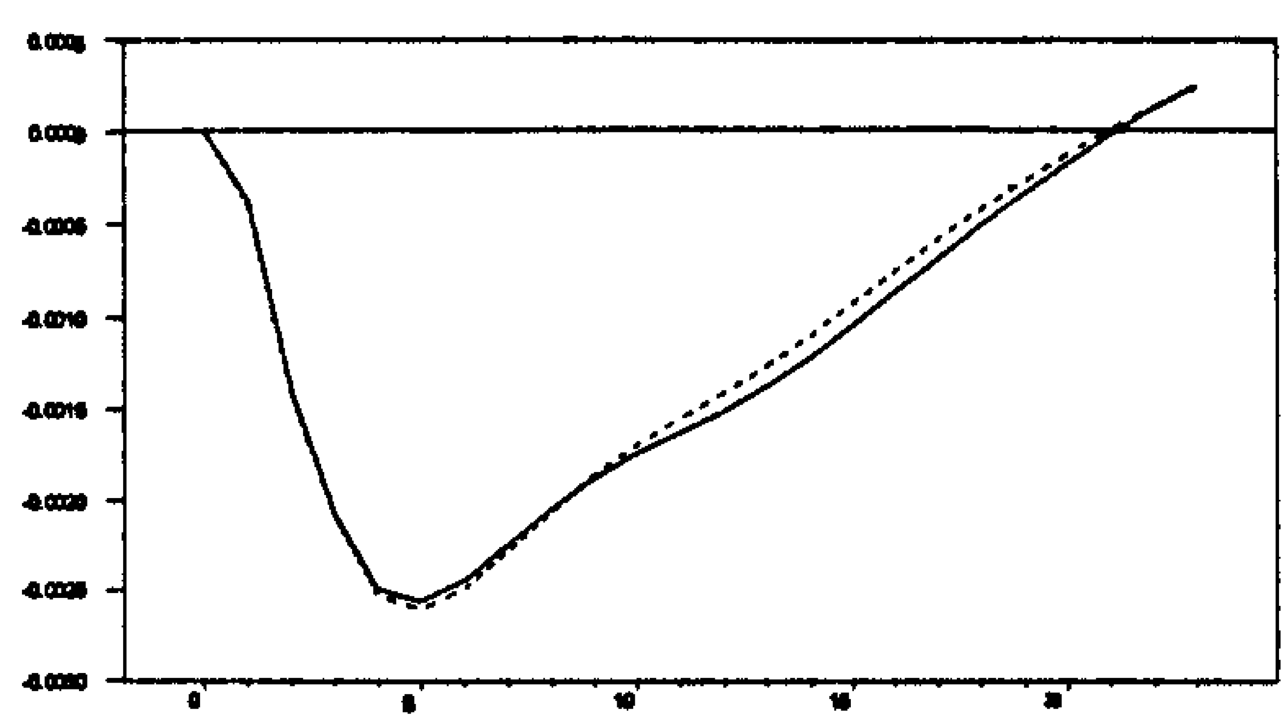
UNITED KINGDOM



NETHERLANDS



SPAIN



SWEDEN

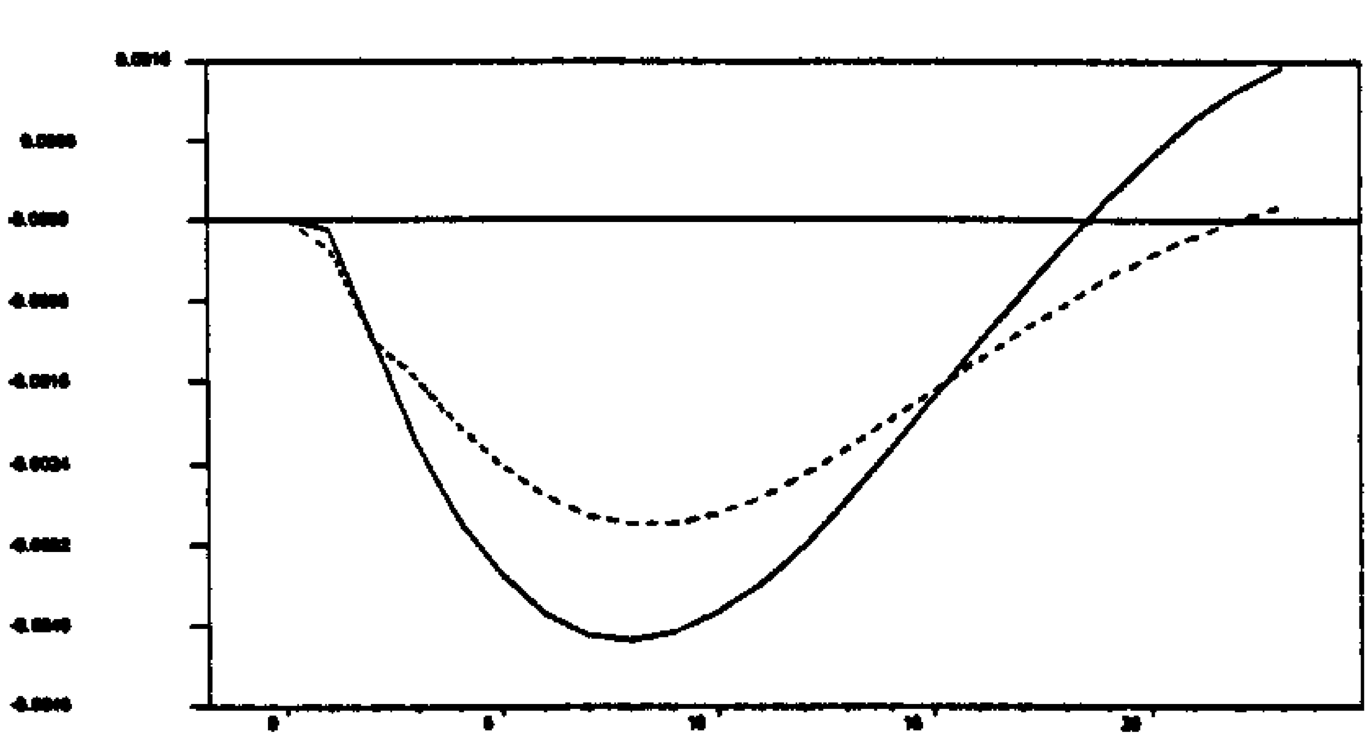
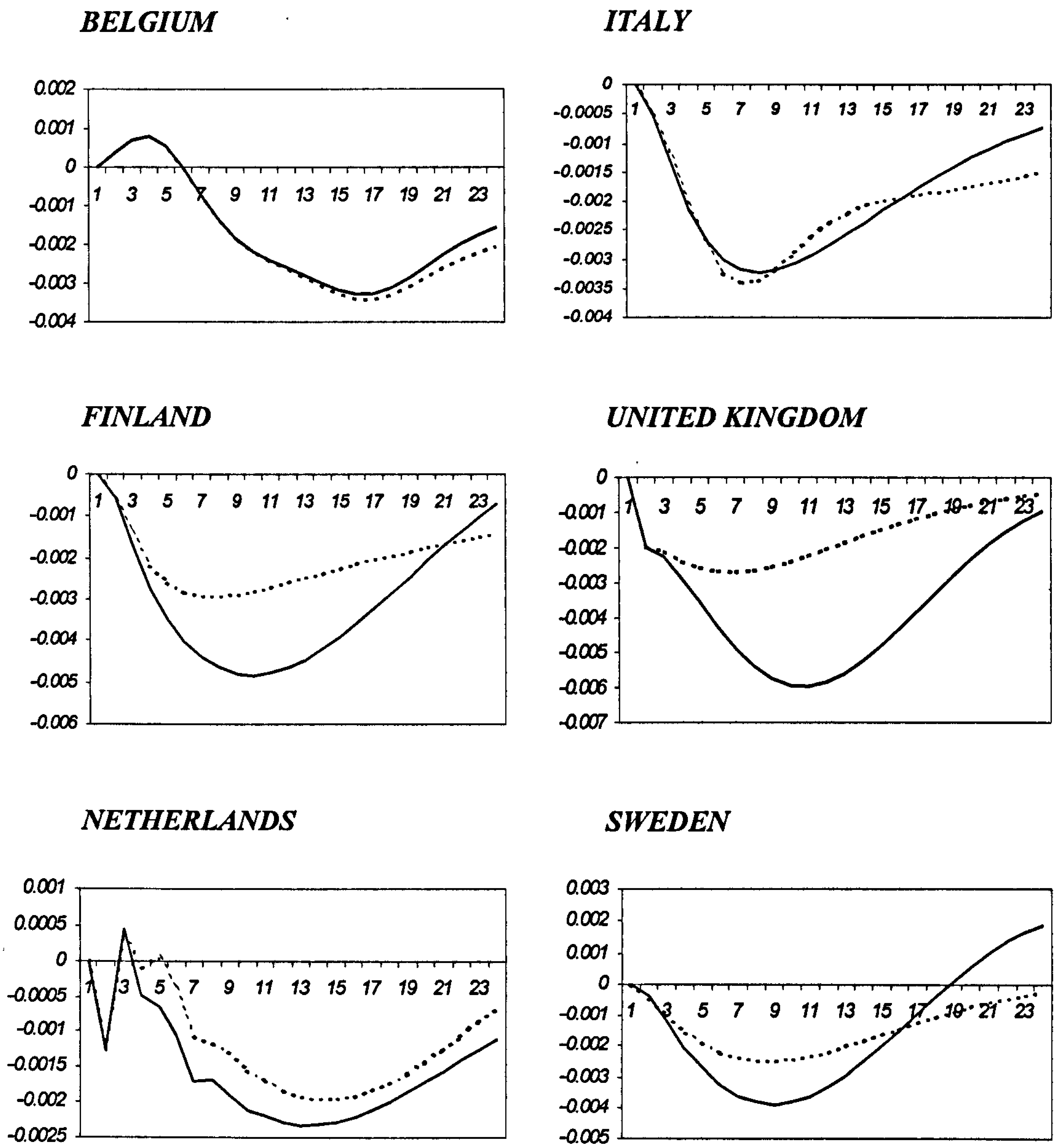


FIGURE A.2.6 **Response of Real Private Consumption to Monetary Shock – Endogenous versus Exogenous Real House Price Models**
(solid line: house price endogenous, solid line: house price exogenous)



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CHAPTER 3

HOUSE PRICE DYNAMICS IN EUROPE: A CROSS-COUNTRY EMPIRICAL ANALYSIS

1. Introduction

Over the last few years, economists and policy makers have shown a renewed interest in the macroeconomic effects of house price movements. In this regard several empirical studies have provided strong evidence in favour of a significant role of residential price fluctuations both for households' consumption decisions or saving behaviour (Kennedy and Andersen, 1994; Girouard and Blöndal, 2001; Ludwing and Sloek, 2002 and Case *et al.,* 2002) and for future movements of the consumer price inflation (Goodhart and Hofmann, 2000).[1] Moreover, changes in house prices have important implications for financial stability (ECB, 2000) and the supply side of the economy (e.g. labour mobility).

[1] Goodhart and Hofmann (2000) examine whether the predictive power of a reduced form equation for inflation can be improved upon by adding asset price variables. They find that housing price movements provide useful information on future inflation and that their predictive power is superior to equity prices and the yield spread.

It is therefore very important for policy makers not only to study the role of residential assets in the monetary transmission mechanism, but also to understand the main determinants of house price fluctuations. Up to now, almost all the empirical evidence has come from the housing literature of the United Kingdom and Nordic economies. Given the crucial role that residential assets play in the macro-economy, recently policy makers called for more effort to be aimed at studying the international house price determinants by using a common methodology and specifications (ECB, 2003). In this regard, the only cross-country studies currently available are very limited (Kennedy and Andersen, 1994; Englund and Ioannides, 1997 and Kasparova and White 2001) and are subject to a number of drawbacks which this chapter aims to overcome. For instance, given the availability of annual frequency time series and relatively short sample periods, some of these papers apply panel data techniques imposing both the same specification and equality of the parameters. Consistent with the presence of heterogeneous mortgage and housing systems across countries, the latter assumption seems very restrictive and throws obvious doubts on the interpretation and the reliability of the estimated pooled coefficients. At the same time, individual single equation results based on annual data are not very satisfactory, probably due to the low number of observations.[2]

Keeping the above limitations in mind, this chapter provides new empirical evidence on the macroeconomic factors driving residential asset values in the EU countries. As in the previous cross-country housing literature, the identification of these forces is first achieved through the implementation of system and panel estimation techniques where the same model specification and equality of the parameters are

[2] The estimated models provided in Kennedy and Andersen (1994) and Englund and Ioannides (1997), for instance, refer to the sample period 1970-1992.

imposed. After formally testing for and rejecting the latter assumption, the chapter makes use of a recent database, containing quarterly information on housing and mortgage markets over the last two decades, and proceeds by estimating separate house price equations within an error correction mechanism (ECM) framework. In allowing for country-specific house price specifications, it emerges that there is a high degree of heterogeneity in the role of the main macroeconomic factors driving house price in the EU countries, which can help to explain the different dynamics of residential asset over the last few decades.

The structure of this chapter is the following. Section 2 provides an overview of house price dynamics over the last two decades and focuses on those financial variables (namely, mortgage and interest rates, and housing lending dynamics) and policy changes in the mortgage markets which, more than others, might have affected the EU housing systems during the sample period. Section 3 then briefly outlines the two main approaches used to model house prices, which will motivate the estimated specifications. Section 4 reviews the previous empirical cross-country literature and discusses its main drawbacks, while section 5 contains the empirical estimation results. In the first subsection, the international housing literature is extended by formally testing for the appropriateness of imposing the same parameter values through the implementation of Seemingly Unrelated Regression (SUR) and dynamic heterogeneous panel estimation methods as suggested by Pesaran, *et al.* (1999a). Given the high degree of heterogeneity, the following subsections proceed by estimating separate country-specific house price equations within the traditional unrestricted error correction mechanism (ECM) framework. After discussing the main results, section 6 concludes.

2. *House Prices and Financial Variable Dynamics: Some Stylised Facts*

2.1 *House Price Developments Since 1980*

Before examining some of the structural factors which are likely to have affected house price developments in the EU countries, this section provides some stylised facts on their fluctuations over the last two decades. The ideal house price index should refer to residential property prices, represent the whole country and cover a consistent part of residential real estate transactions in each period. Moreover, it should take changes in the quality of the property into account. Such information, however, is not available for all the EU countries. While house price indices in the U.K., Ireland, Finland and Sweden are good, national time series data for most of the other European countries are less satisfactory.[3] Appendix 1 provides a full description of the database used in this chapter, which, in its current form, represents the best available official information on housing markets in the EU area.[4]

Figure 3.1 shows the annual real house prices (deflated by the harmonised consumer price index - HICP) for the EU countries over the period 1980-2001.[5] The countries in Panel A (Denmark, Finland, Italy, Sweden and the United Kingdom) exhibit a bell-shape, with different amplitude and timing of the peaks. Whereas Austria, Germany, Bèlgium, Luxembourg and, to some extent, France and Portugal do not display any particular cyclical patter, Ireland and the Netherlands have experienced a

[3] Cross-country comparability and quality of national house price indices are not the only limitations. Indeed, housing markets are rather heterogeneous within the same country. House price dynamics are very much related to regional supply-demand factors, which might generate different house price cycles. As pointed out by several authors (Kasparova and White, 2001), however, regional differences might be smaller than differences across countries due to the fact that financial, institutional and legal systems are common to all regions within the same national border. Meen and Andrew (1998a) provide an excellent review of the literature on modelling regional house prices.

[4] The database used in most of the international housing modelling empirical literature is provided by the Bank for International Settlements (BIS) and publicly available on an annual basis only. In our data set, an effort has been made to collect the same series, but at higher frequency.

[5] Greece is not included because the house price index starts in 1995 only.

steady increase in residential price levels since the mid-1980s, with a significant acceleration starting from the mid-1990s.

Table 3.1 provides the standard deviation, the cumulative growth and the coefficient of variation of the annual real house prices over the same period. With the only exception of Germany, the overall cumulative growth during the period 1980-2001 was positive in all countries. Ireland, Luxembourg, the Netherlands, Spain and the U.K. show the highest figures (between 76% and 87%), while, on the other side of the spectrum, there is Belgium, Denmark, France, Italy, Portugal and Sweden with cumulative growth rates smaller than 35%. Similar to the results obtained in Kennedy and Andersen (1994), who use Bank for International Settlements (BIS) annual data for the period 1970-92, nominal house prices tend to rise at a higher pace than the consumer price index in the long run.[6] With the exception of a few cases, the table shows significant sub-sample differences not only across, but also within the same country during the 1980s and the 1990s.

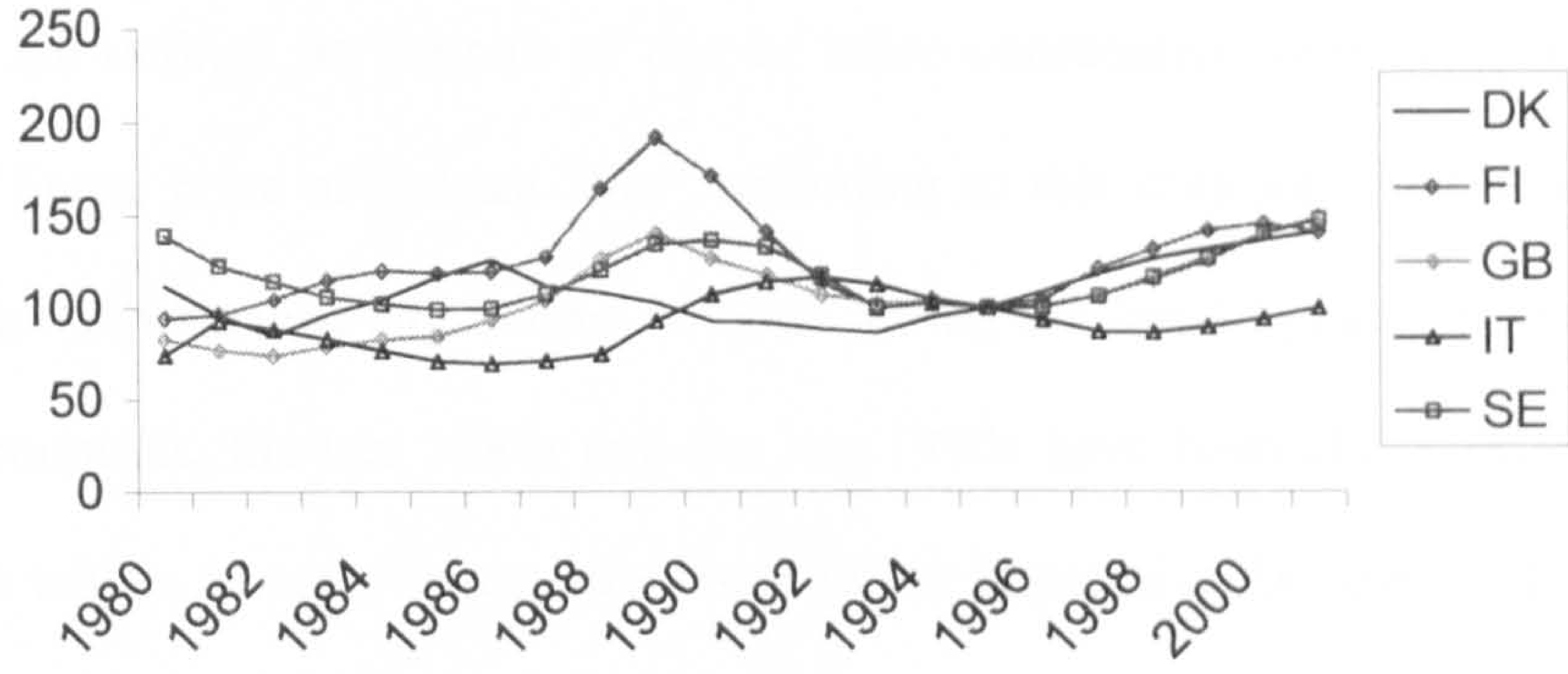
In Finland, France, the Netherlands, the U.K., Luxembourg and Ireland, price volatility seems to be higher. Slightly lower figures are found in all the remaining countries. By comparison to the rest of Europe, real house prices in Germany and Portugal are less volatile. This rough classification is maintained once the coefficient of variation is taken into account. Differences in the volatility of house price across countries might be related to a number of specific demand and supply factors. Many authors have emphasised the role of the financial liberalisation process in the U.K. and other Nordic countries, and argued that, as a result of the deregulation, the above

[6] In the overlapping sample period, the BIS house price indexes coincide with our data. The only exception is Germany, where different indicators are available. BIS data set does not include Luxembourg, Portugal and Austria.

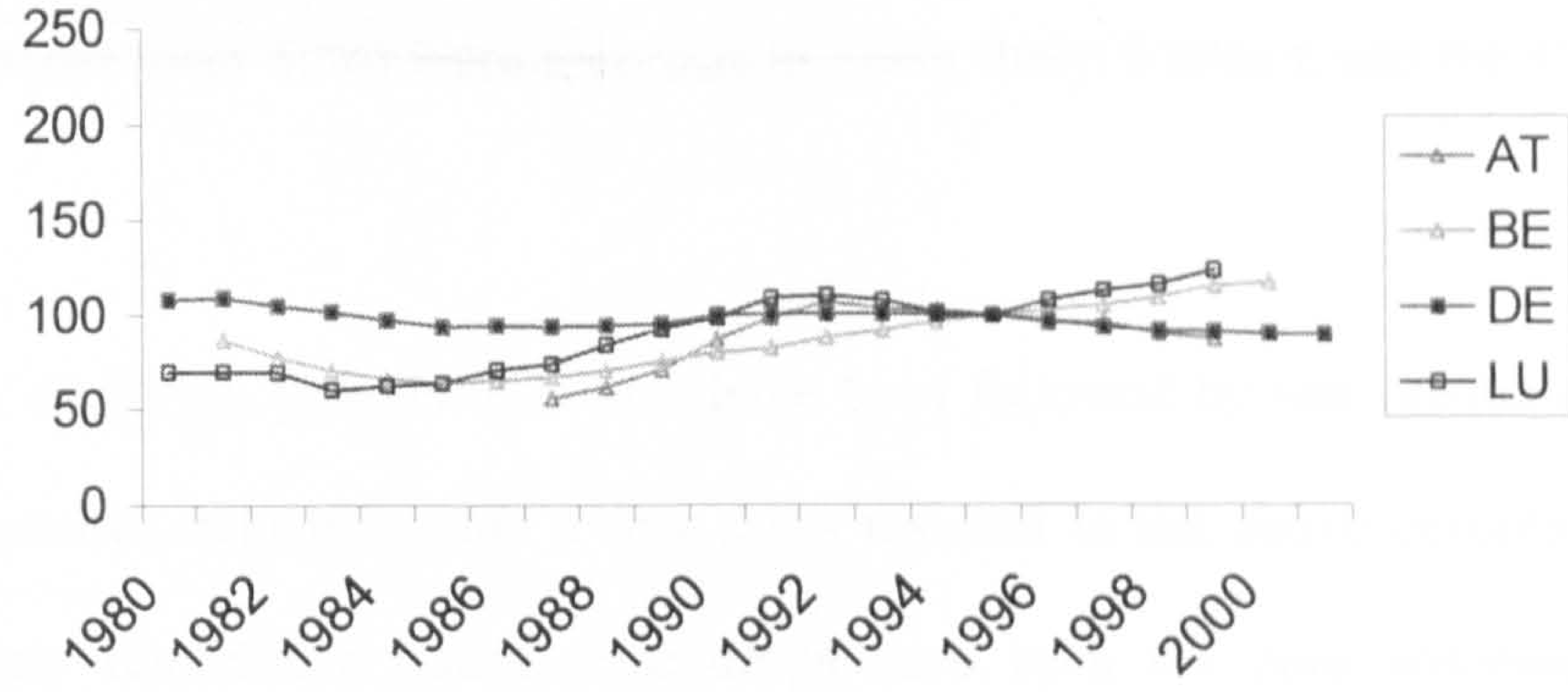
countries witnessed higher volatility (European Commission, 1999). Although Finland, Sweden and the U.K. seem to belong to the 'high-volatile' house price group, however, they do not stand out from other countries, in which the mortgage market went through less intense changes. Expressing volatility as the standard error of the regression of real house prices on a constant and a time trend, the pattern is more in line with the above deregulation argument. Finland, Ireland, the Netherlands, Sweden and the U.K. had relatively more volatile prices over the period 1980-2001, whereas France, Belgium, Luxembourg, Germany and Portugal show more stable residential prices.

FIGURE 3.1 Real House Prices in the EU (Index 1995=100)

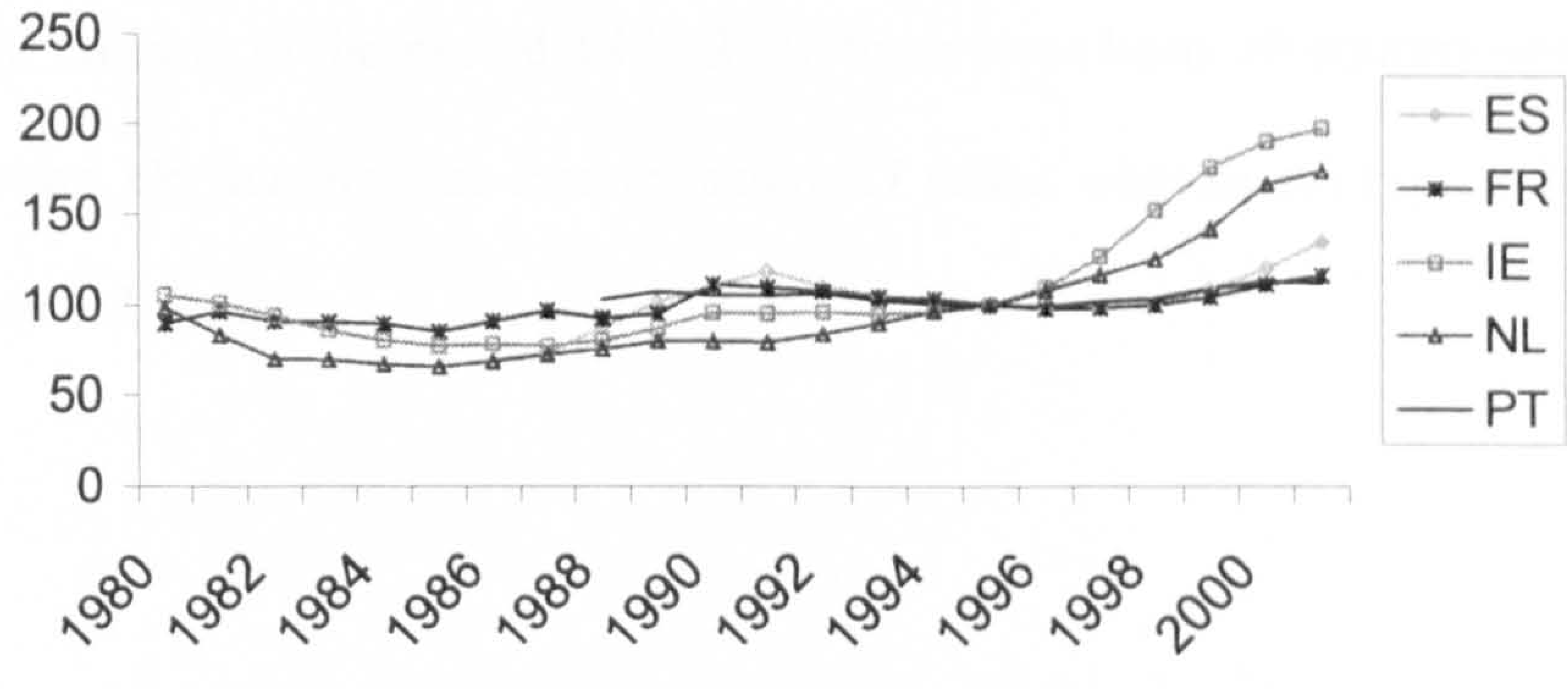
Panel A



Panel B



Panel C



Sources: See Appendix 1 for a description of the series. House prices in Greece are only available from 1995. As a result, it is not include in the sample. AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, PT = Portugal, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom, LU = Luxembourg.

Residential house price dynamics can be further investigated by looking at the booms and busts, that is, periods of sharp growth and decline of real house prices. Although it is not clear what the correct or 'official' definition of a boom or bust is, in Table 3.2 they are defined as periods of one or more consecutive years with yearly changes in real house price of at least 8%.[7] According to this criterion, Germany and Portugal have not experienced any booms or busts during the last 20 years. In most of the remaining countries, the late 1980s and the late 1990s have been characterised by strong increases which, in some cases, have been rather impressive. In Austria, Ireland and France, for instance, the cumulative real house price growth reached 92%, 90% and 87% during the period 1988-92, 1996-2000 and 1987-90, respectively. During the late 1980s strong booms (over 50%) were recorded in Spain, Italy, Finland, and the United Kingdom.

In many countries, periods of booms have been followed by less dramatic real house price decreases, which in only a few cases resulted in the above definition of busts. Particularly remarkable house price drops have been the ones witnessed in Finland, Sweden, France and the United Kingdom (respectively, 48%, 25%, 22% and 19%) during the early 1990s. Overall, in the European countries, excluding Luxembourg and Greece, in the period 1980-2001 there have been 19 country-specific booms and 14 busts. On average, the former lasted 2.7 years, whereas the typical length of a bust has been 1.8 years.

[7] The 8% threshold is arbitrary, but very close (although more conservative) to the criterion used in ECB (2003). Here, booms and busts are defined as periods with uninterrupted changes in real house prices of at least 10% annually.

TABLE 3.1 Growth and Volatility of Real House Prices in the EU

	1980-2001	1980-1990	1990-2001		1980-2001	1980-1990	1990-2001
Austria				***Belgium***			
Standard deviation	16.3	7.7	6.7		17.5	7.4	12.5
Cumulative growth	56%	56%	0%		35%	-7%	46%
Coefficient of variation	18.3%	12.2%	6.9%		20.1%	10.3%	12.6%
S.E.	13.6**	-	-		8.8**	-	-
Denmark				***Finland***			
Standard deviation	16.7	11.8	20.2		25.8	30.5	22.6
Cumulative growth	26%	-17%	52%		50%	82%	-17%
Coefficient of variation	15.5%	11.2%	18.5%		20.5%	24.4%	17.9%
S.E.	14.9**	-	-		25.6	-	-
France				***Germany***			
Standard deviation	23.5	19.2	19.8		5.4	6.0	4.6
Cumulative growth	29%	23%	5%		-17%	-7%	-10%
Coefficient of variation	24.5%	23.8%	18.4%		5.5%	6.0%	4.8%
S.E.	5.8**	-	-		4**	-	-
Ireland				***Italy***			
Standard deviation	36.6	10.2	40.6		15.3	9.3	10.5
Cumulative growth	87%	-9%	107%		35%	44%	-6%
Coefficient of variation	33.5%	11.8%	31.9%		17.1%	12.0%	10.4%
S.E.	25.2**	-	-		13.2**	-	-
Luxembourg				***Netherlands***			
Standard deviation	20.9	10.1	7.8		31.3	10.0	32.7
Cumulative growth	77%	41%	25%		76%	-19%	117%
Coefficient of variation	23.2%	14.0%	7.2%		32.5%	13.2%	28.8%
S.E.	7.9**	-	-		19.1**	-	-
Portugal				***Spain***			
Standard deviation	4.5	3.1	4.8		14.3	13.9	11.2
Cumulative growth	10%	2%	8%		83%	49%	23%
Coefficient of variation	4.3%	2.9%	4.5%		13.6%	15.8%	10.3%
S.E.	4.2**	-	-		11.9**	-	-
Sweden				***U.K.***			
Standard deviation	15.9	14.4	17.4		21.8	22.4	16.2
Cumulative growth	6%	-2%	8%		79%	52%	17%
Coefficient of variation	13.6%	12.6%	14.6%		20.6%	23.8%	13.9%
S.E.	16.1	-	-		15.3**	-	-

Notes: Austrian data is limited to the period 1987-99. Luxembourg 1980-1999. For Belgium, Spain and Portugal the sample starts in 1981, 1987 and 1988, respectively. The coefficient of variation is defined as the ratio of the standard deviation over the mean. S.E. is the standard error of the regression of annual real house price index on a constant and a time trend over the period 1980-2001. (**) indicates 5% statistical significance of the trend.

TABLE 3.2 House Price Boom and Busts over the Period 1980-2001

	Periods of Booms (Busts)	*Cumulative Growth (Decrease)*
Austria	1988-1992	+91.3
Belgium	(1982-1983)	(-18.8)
Denmark	(1980-1982)	(-33.6)
	1983-1985	+37.7
	(1987)	(-11.1)
	(1990)	(-9.9)
	1994	+9.9
	1996-1997	+18.33
Finland	1982-1983	+19.4
	1988-1989	+50.1
	(1990-1993)	(-47.8)
	1997-1998	+26.2
France	1987-1990	+86.9
	(1992-1993)	(-21.6)
	(1996)	(-10.2)
U.K.	1986-1989	+65.9
	(1990)	(-9.7)
	(1992)	(-9.3)
	1999-2000	+21.9
Ireland	(1983)	(-8.8)
	1990	+9.6
	1996-2000	+90.5
Italy	1981	+25.6
	1988-1992	+75.6
	(1994)	(-9.85)
Luxembourg	(1983)	(-13.2)
	1986	+10.5
	1988-1989	+25.32
	1991	+10.7
	1996	+8.1
Netherlands	(1980-1982)	(-37.7)
	1997	+8.2
	1999-2000	+32.8
Spain	1988-1990	+49.2
Sweden	(1980-1981)	(-21.1)
	1987-1989	+34.9
	(1992-1993)	(-24.9)
	1998-2000	+31.8

Notes: Booms (busts) are defined as periods of one or more consecutive years with yearly increase (decrease) in real house price of at least 8%. According to this criterion Portugal and Germany did not experience any boom or burst during the period 1980-200. All figures are in percentage. Greece is not included.

2.2 *Financial Liberalisation and the Mortgage Market*

Many authors argue that the deregulation process of the banking and financial markets affected the functioning of the housing market not only during the liberalisation period itself, but also in the longer term. This might imply that agents' (households and credit institutions) behaviour might be more sensitive to shocks in those countries where housing finance is more developed.[8]

Until the beginning of the 1980s, the mortgage markets of most EU countries were strictly regulated by national authorities. Such regulations were related to lending and deposit interest rate ceilings, portfolio restrictions for financial institutions, limits on loan-to-value ratios and repayments periods and quantitative credit controls, which all reduced access to mortgage lending. Moreover, the market was dominated by few specialised institutions, which limited the entry of other banks or financial institutions into the housing mortgage market, reducing the competition in the market for new mortgages. With the only exception of Germany, where the financial system was relatively mature already by the end of the 1970s, in most countries significant financial reforms were introduced, starting in the early 1980s (Lomax, 1994).

Table 3.3 provides a brief overview of the main policy measures that were introduced during the 1980s and 1990s. They included reduction in the barriers to entry to the housing finance industry, relaxation and subsequent abolition of interest rate ceilings and credit controls, relaxation of portfolio restrictions for financial institutions, abolition of exchange rate controls and the introduction of securitisation. The above

[8] Meen (1996) and Maclennan *et al.* (1998) argue that in countries with more developed mortgage systems, agents might react more strongly to house price and interest rate changes. Iacoviello and Minetti (2002) provide some evidence in this direction. More recently, Lamont and Stein (1999) and Almeida (2001) show a higher sensitivity of house prices to specific shocks (i.e. income shocks) where homeowners are more leveraged and financial constraints less binding.

developments hugely increased competition in the mortgage market, reducing quantitative rationing and expanding the range of mortgage products available. While in the United Kingdom and Nordic countries this process was relatively quick, many continental European economies showed a slower pace of deregulation, which partly explains the current differences in the features of the mortgage market and different dynamics of the house price.

Figures 3.2 and 3.3 show the ratio of the stock of housing lending over GDP (MSTOCKY) and its growth over the period 1980-2001. The countries are split according to the periods in which MSTOCKY shows similar dynamics. For instance, in Denmark, Finland, Ireland, Sweden and the U.K., the financial deregulation measures seem to have had most of their effects in the 1980s with growth rates over 10-15% during the period 1982-89. In particular, the U.K. displays a pronounced growth pattern since the early 1980s, with Denmark, Sweden and Finland presenting similar rates in the subsequent years.

The early 1990s are characterised by a relatively stable and (with the only exception of Ireland) decreasing level of housing lending over GDP which started growing just over the most recent period. Spain presents a pattern similar to the above countries, that is, a strong credit expansion during the second half of the 1980s, which continued to grow over the following years. At the other side of the spectrum, there are Italy, Greece and Portugal, where housing lending accelerated mainly during the last decade, with growth rates reaching figures greater than 15%. The Netherlands presents a persistent increase in housing lending, although the rates have generally been smaller than 10%. The remaining countries, namely Belgium, Germany and France, experienced a relatively stable lending volume over the period under investigation.

Although it is clear that such growth figures have to be interpreted with the level of the outstanding stock of mortgage lending over GDP - which differs quite significantly from country to country[9] - these developments in the mortgage market might explain part of the housing dynamics discussed above. For instance, strong credit growth patterns of the 1980s in the U.K., Sweden, Finland and Denmark are associated with significant housing booms in the same period. A positive relationship seems to emerge also from other European countries in other periods. The Netherlands and Portugal during the 1990s are clear examples.

However, given that house prices and the value of the mortgage stock are contemporaneously determined, it is not easy to evaluate the effective role of mortgage market dynamics on housing prices from a simple visual inspection or unrestricted correlation analysis. Rising house prices are probably an important reason for mortgage debt to have increased so sharply during specific periods, but, as emphasised in ECB (2003), there are other factors which greatly contributed to the accumulation of mortgage debt. Amongst them, there are improving income expectations, falling mortgage interest rates and favourable fiscal treatment of housing lending. Section 4.2 of this chapter will discuss how the empirical literature accounts for the role of the mortgage market deregulation (and dynamics) in house price studies, and how these effects are tested for in the empirical section.

[9] Relatively high ratios are present in Denmark, Germany, the Netherlands, Sweden and the United Kingdom, whereas the lowest figures are in Greece and Italy (see the previous chapter and Table 3.4 in section 2.4 below)

TABLE 3.3 Policy Changes in the EU Affecting the Mortgage Market

Austria	Liberalisation of interest rates in 1980 Abolition of credit controls in 1981 Re-establishments of interest rate controls through interest rate cartels in 1985 Prudential reforms, capital requirements tightened in 1987 Interest rate cartels expires 1993 Saving bank reforms in 1999 Privatisation of state-owned banks in 1992-2000
Belgium	Introduction of variable interest rate loans in 1992 (amended in 1995) Wave of mergers and privatisation in the banking sector in the 1990s
Denmark	Liberalisation of mortgage contract terms in 1982 Interest rate deregulation in 1982 Elimination of restrictions on mortgage bond issuance in 1989
Finland	Funding quotas from the Central Bank to commercial banks eliminated in 1984 Abolition of interest rate controls in 1986 Government withdrew guidelines on mortgage lending in 1987 Securitisation introduced in 1989
France	Bank specialisation requirements reduced in 1984 Elimination of credit controls 1987 Reform of securitisation of mortgage loans in 1999 Reduced limits on early repayments fees
Germany	Interest rate deregulation in the 1970s
Greece	Gradual liberalisation of quantitative constraints, interest rates and other terms and conditions on housing loans from mid-1980s to early 1990s Liberalisation of mortgage refinancing and expansion of non-specialised commercial banks into mortgage lending in the late 1990s
Ireland	Formal guidelines for bank lending to private sector ended in 1984 Interest rate deregulation in 1985 Relaxation of exchange controls in 1988 Reductions in the primary liquidity ratio from 8% to 2% during the period 1991-99 Securitisation introduced in the second half of 1990s

Sources: Girouard and Blondal (2001) (Table 3, page 23) and ECB (2003) (Table A.3, page 54).

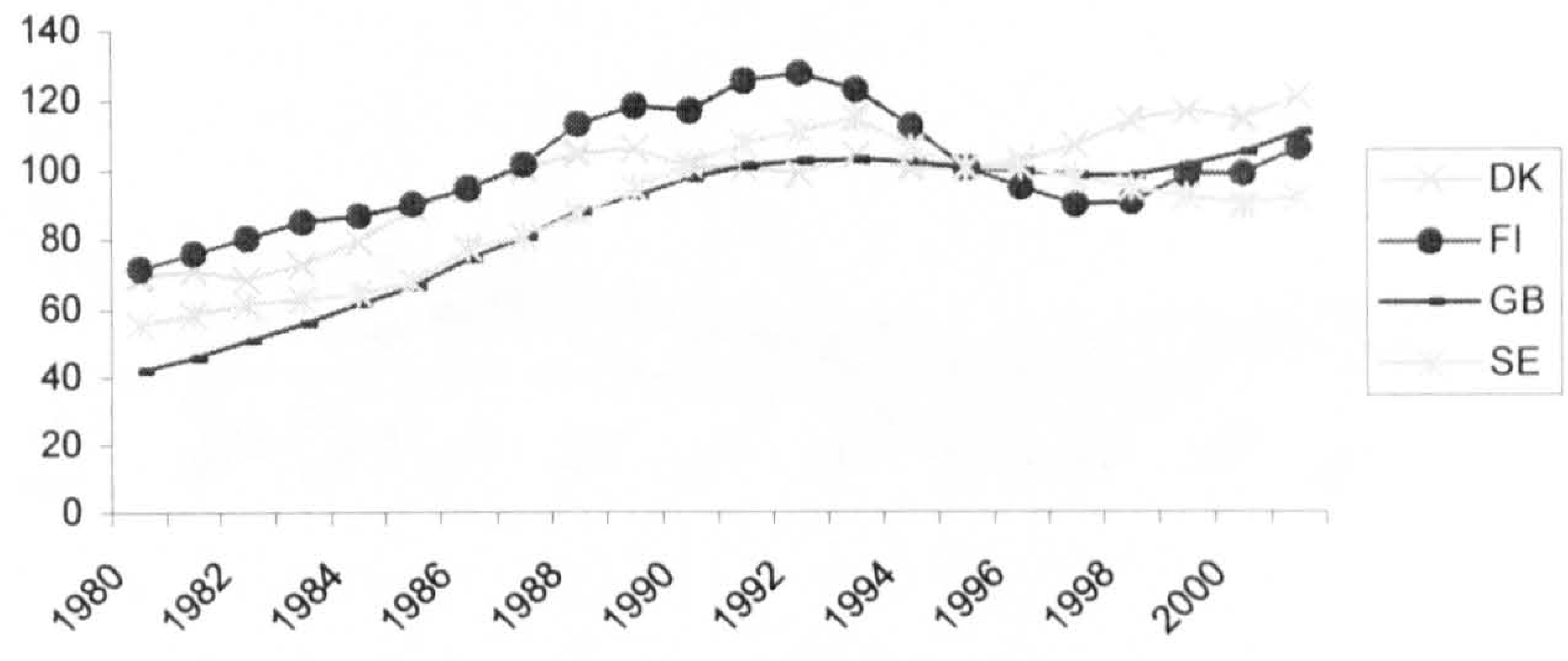
TABLE 3.3 Policy Changes in the EU Affecting the Mortgage Market (continued)

Italy	Interest rate deregulation in 1983 Credit ceilings eliminated in 1983 and temporarily re-imposed in 1986 and 1987 Abolition of administrative controls on branching in 1990 Separation of long-term and short-term credit institutions abolished in 1993 Increased of legally maximum LTV from 75% to 80% in 1995
Luxembourg	None
Netherlands	Interest rate deregulation in 1980 Relaxation of the lending criteria since 1992
Portugal	Easing of entry restrictions in the banking and insurance sector from 1983 Liberalisation of interest rates (in 1984 for deposit rates and 1989 for lending rates) Abolition of credit controls and credit guidelines in 1990 and 1991 Liberalisation of entry, branching, specialisation and segmentation restrictions early 1990s
Spain	Interest rate liberalisation from 1974-81 and in 1987 Abolition of differences in the activities permitted for different types of banks in 1988 Saving banks allowed to open branches outside their home regions in 1988 Securitisation of mortgage loans introduced in 1992 Introduction of upper limits on cancellation fees in 1994 and 1996
Sweden	Mortgage institutions free to issue bonds for refinancing of old dwellings in 1983 Loan ceilings for banks abolished in 1985 Portfolio regulations on insurance companies dropped in 1986
U. K.	Removal of foreign-exchange and elimination of credit controls ("the corset") in 1980 Bank of England's minimum lending rate abolished in 1981 Banks allowed compete with building societies for housing finance after 1981 Building societies allowed expand their lending business after 1986 Government withdrew guidelines on mortgage lending in 1986 Securitisation introduced in 1987

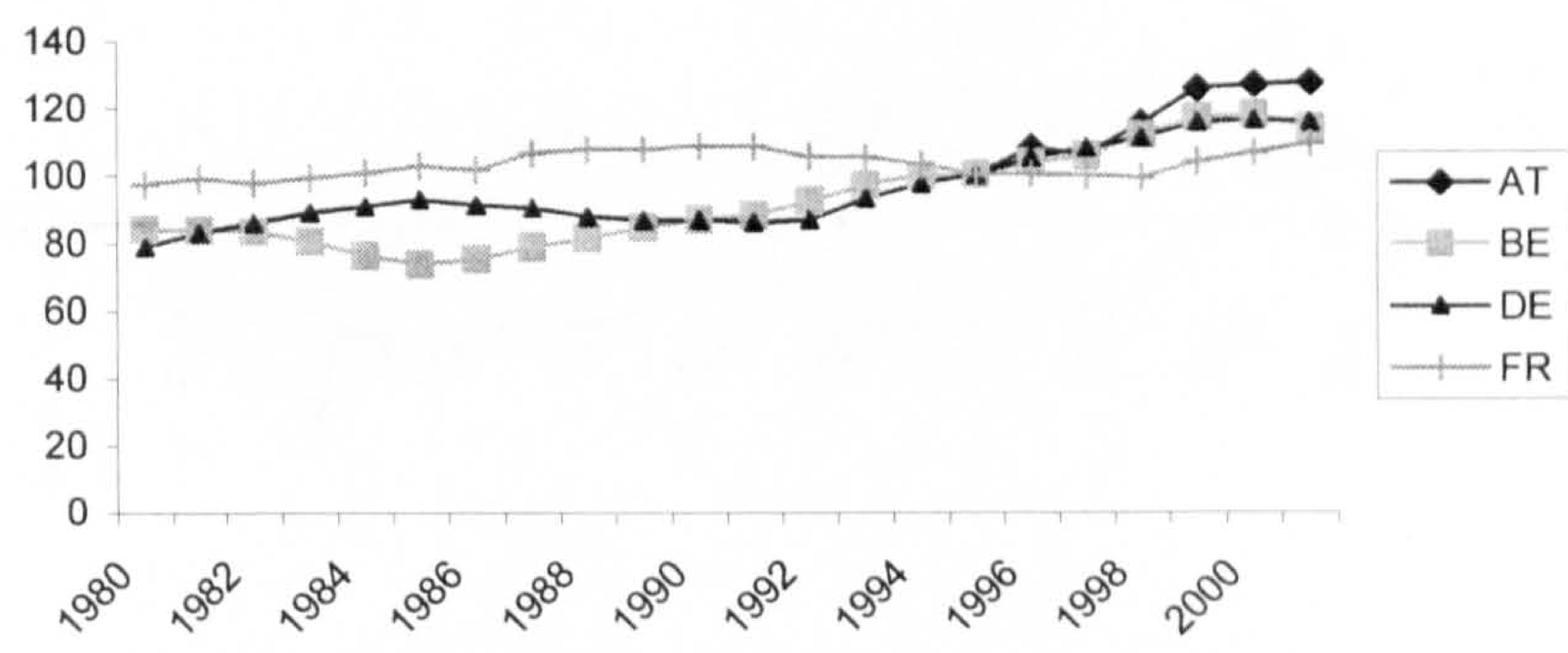
Sources: Girouard and Blondal (2001) (Table 3, page 23) and ECB (2003) (Table A.3, page 54).

FIGURE 3.2 Ratio of Housing Lending Stock over GDP (Index 1995=100)

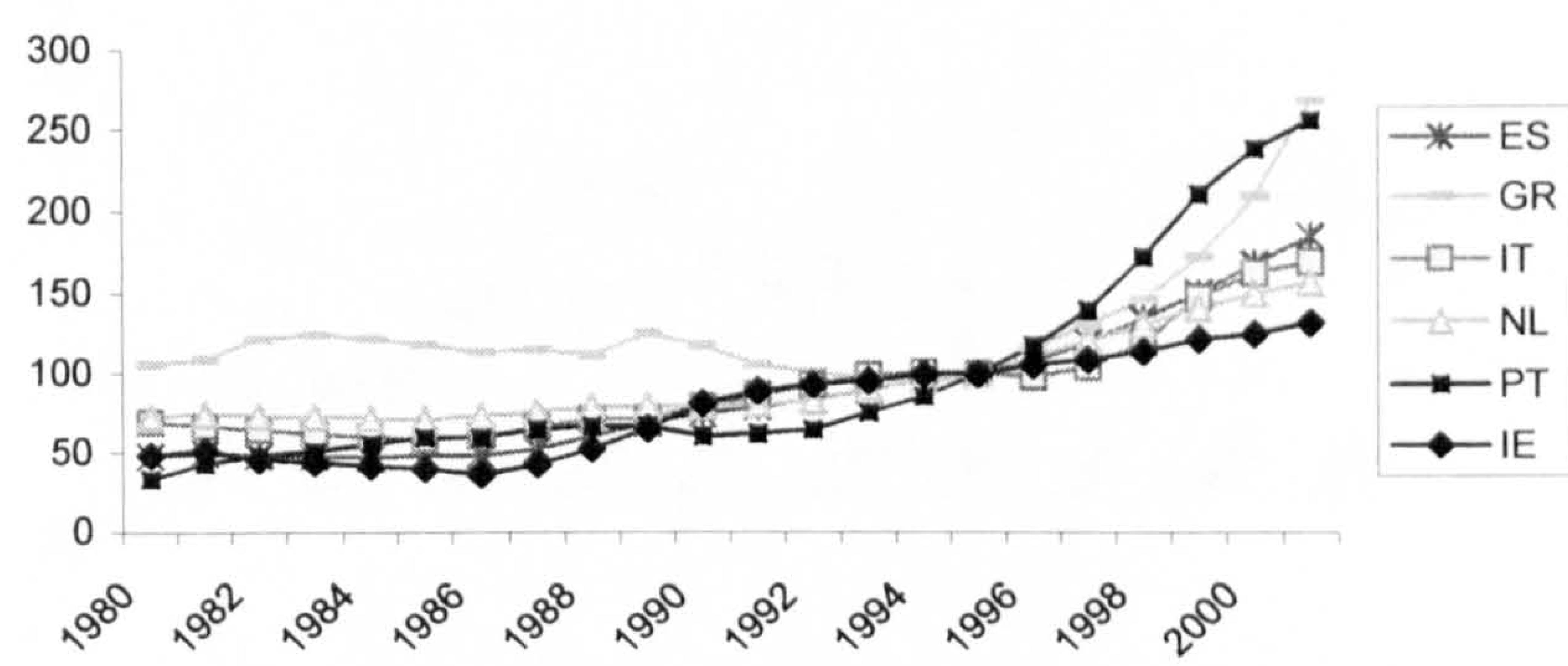
Panel A



Panel B



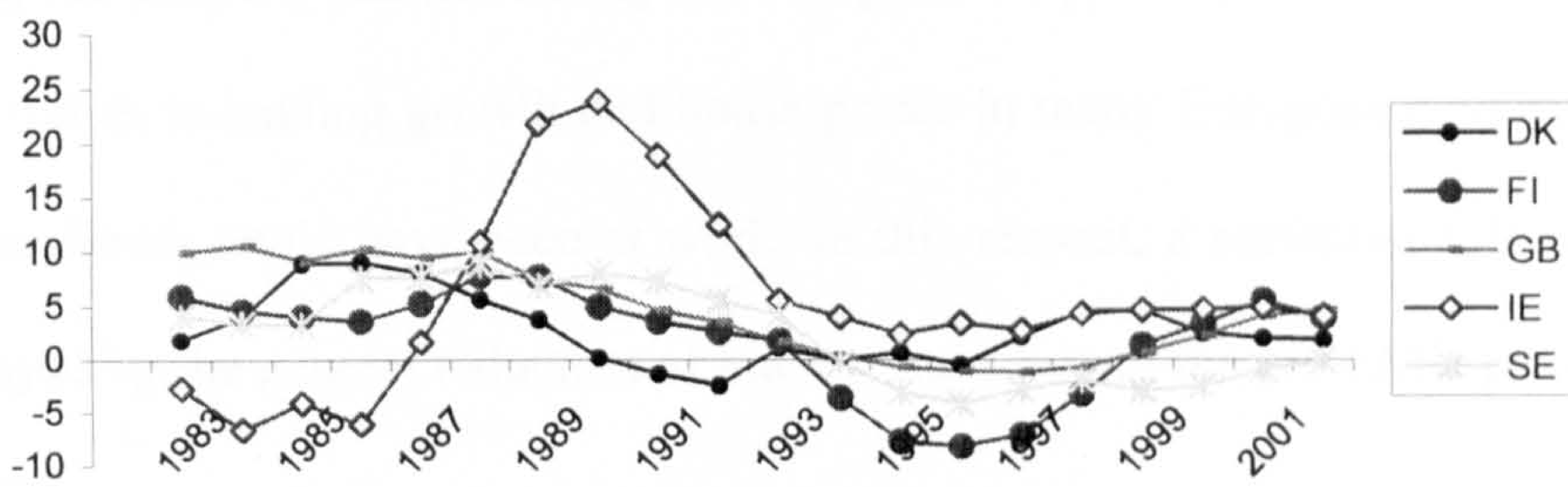
Panel C



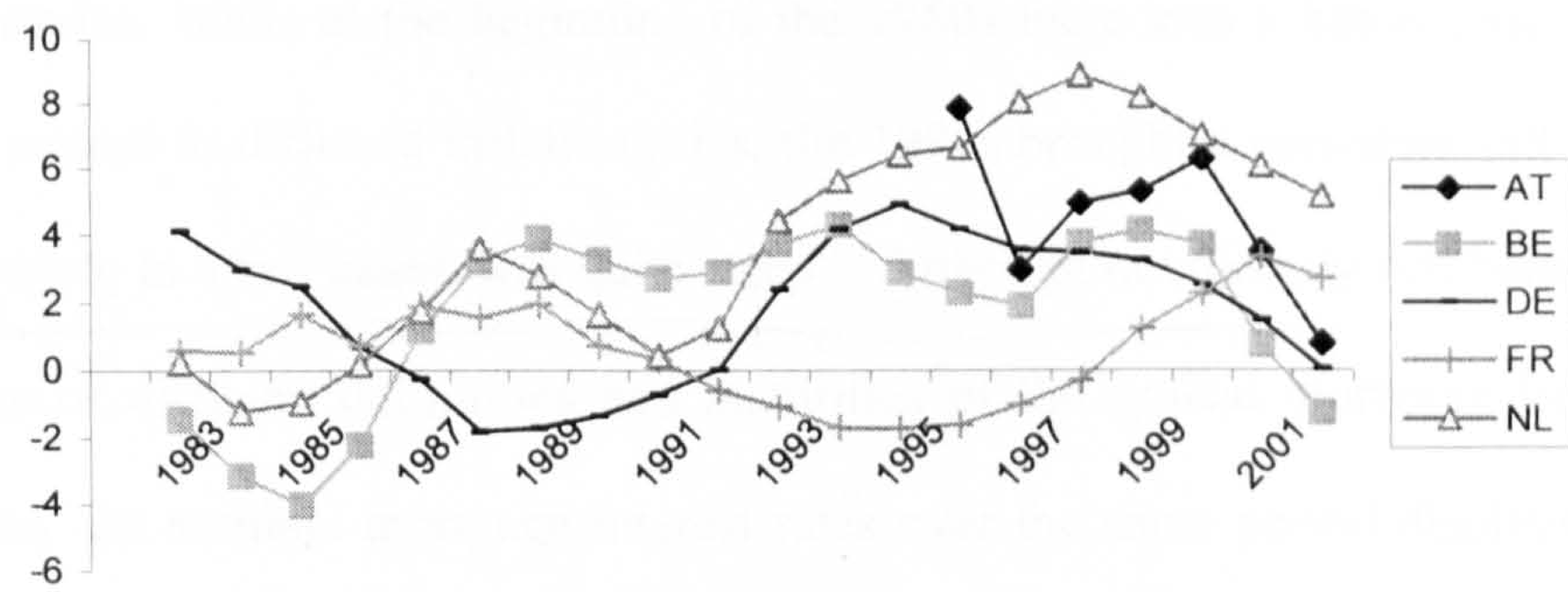
Notes: AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, PT = Portugal, GR = Greece, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom.

FIGURE 3.3 Growth of Housing Lending Stock to GDP Ratio

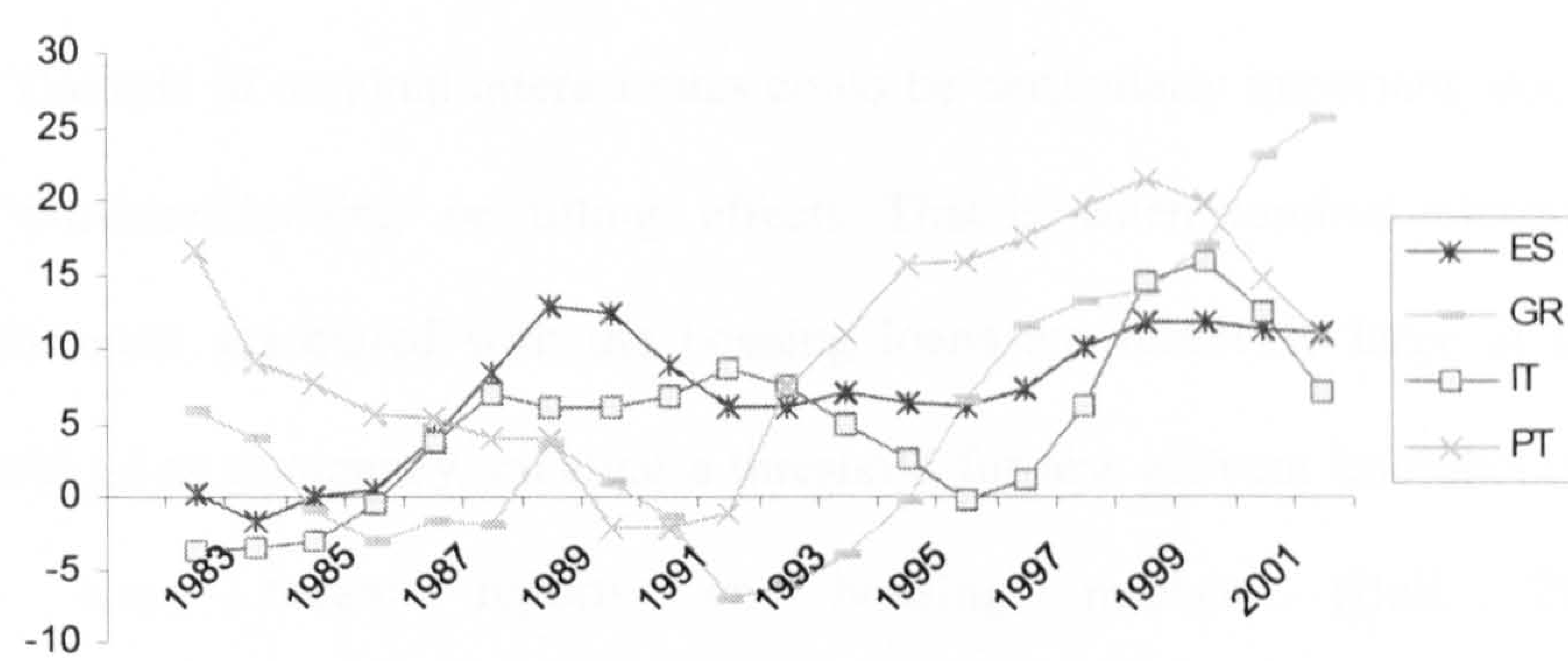
Panel A



Panel B



Panel C



Notes: 3-year moving average (centered on the 2nd year) of the growth rate of the ratio of housing lending to GDP index. AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, PT = Portugal, GR = Greece, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom.

2.3 *Real and Nominal Interest Rates Dynamics*

While developments in credit availability might have been one of the driving factors affecting the property market during the deregulation process of the 1980s, the recent upward trends in lending growth and house prices in many European countries implies that other forces could have been at work. In this respect, a particular role could have been played by the general reduction of interest rates during the pre-EMU period.

Figure 3.4 shows the dynamics of the nominal long-term interest rates for the EU countries. While at the beginning of the 1980s there was a heterogeneous pattern mainly related to different inflation rates, the 1990s brought a persistent fall in interest rates, which, in a few cases, was quite steep.[10] Although not directly comparable for the presence of different definitions and maturities of the typical mortgage loans across countries, the nominal mortgage interest rates over the same period displays a similar pattern (Figure 3.6).[11]

The role of nominal interest rates could be particularly important, due to the so-called 'front-end loading' or 'tilting' effects. That is, when nominal interest rates are high, the costs associated with the housing loans are relatively large at the time a household takes occupancy, creating a threshold for less affluent households.[12] In this regard, some recent reports on housing markets (Ball, 2002 and

[10] Nominal long-term interest rates in Greece and Portugal are available only from 1993 and 1985, respectively. As a result they are not included in the Figure. From the available years, however, the respective rates fell from values greater than 20% to 5%.

[11] A simple comparison based on the *level* of mortgage interest rates should be avoided, because the series refers to different mortgage products and, in some cases (within the same country) have been constructed as a weighted or simple average of different mortgage rates. See the Appendix 1 for further explanations and a description of the series used in this chapter.

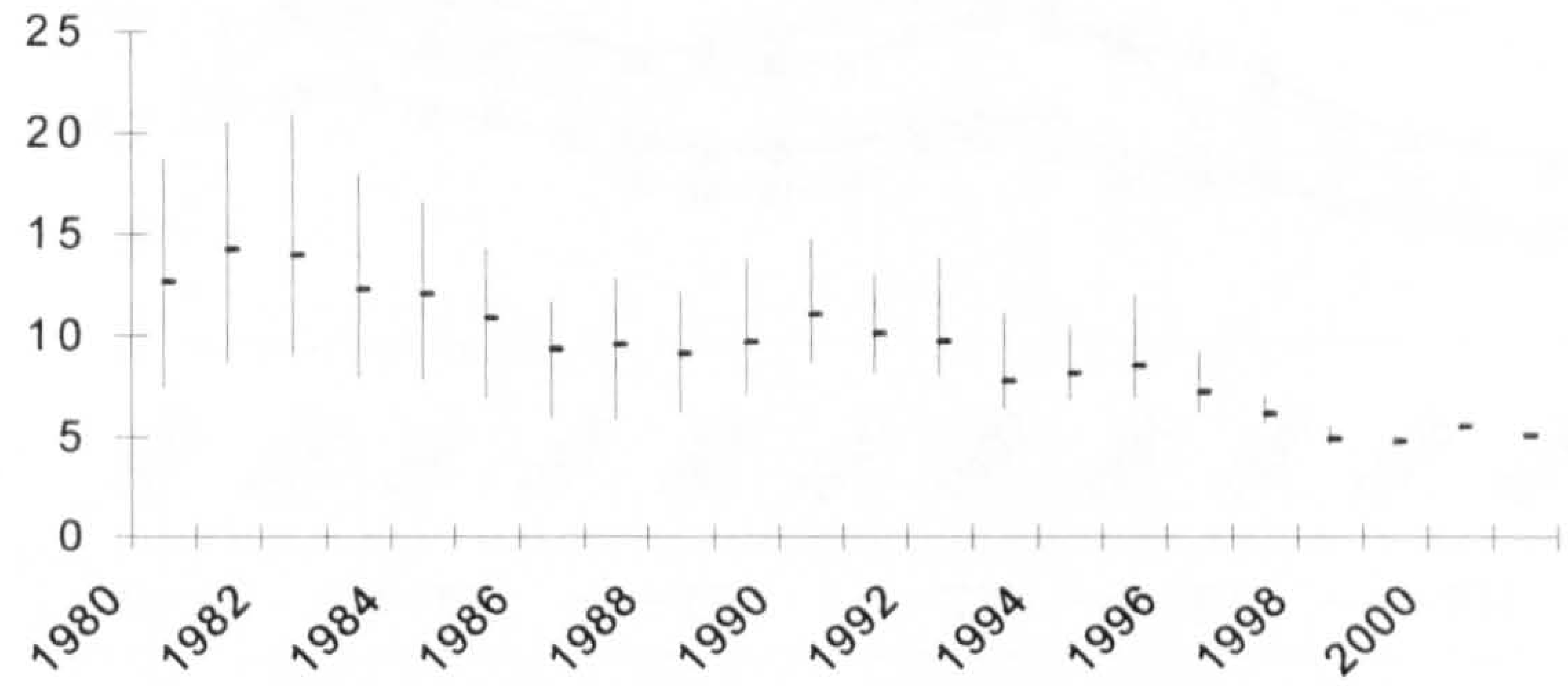
[12] See Colwell and Dehring (1996), Miles (1994) and Bank of England (1991) for a discussion of the tilting effect. Schwab (1981, 1982) and Kearl (1978) provide a very simple example of how an increase in expected inflation (which raises nominal interest rates) does 'tilt' the stream of real payments forward, possibly leading to cash flow problems for some home-buyers. "*Consider a 30-year mortgage with a $20,000 principal which bears a real interest rate of 3%. If expected inflation were zero, the nominal and real annual payment would be constant $1,020. If expected inflation were 8%, however, the real payment would be $2,130 in the first year and $229 in the thirtieth.*"

PricewaterhouseCoopers, 2002) have argued that the unprecedented fall of nominal interest rates could be considered the prime factor explaining the recent booms in housing demand from the mid 1990s.

The actual effects of falling nominal rates on households borrowing decisions could be better appreciated by taking into consideration the offsetting effects of rising mortgage debt. Figure 3.7 provides an approximation of the debt servicing which is calculated as the ratio between the product of nominal mortgage interest rates and the housing nominal debt stock over the nominal household disposable income. From this rough estimate it emerges that the debt servicing of the 1990s has been either relatively stable or following a progressive reduction. The only exceptions are the Netherlands, Portugal, Italy, Ireland and Spain, where an increase can be observed during the last 2-3 years of the sample period.

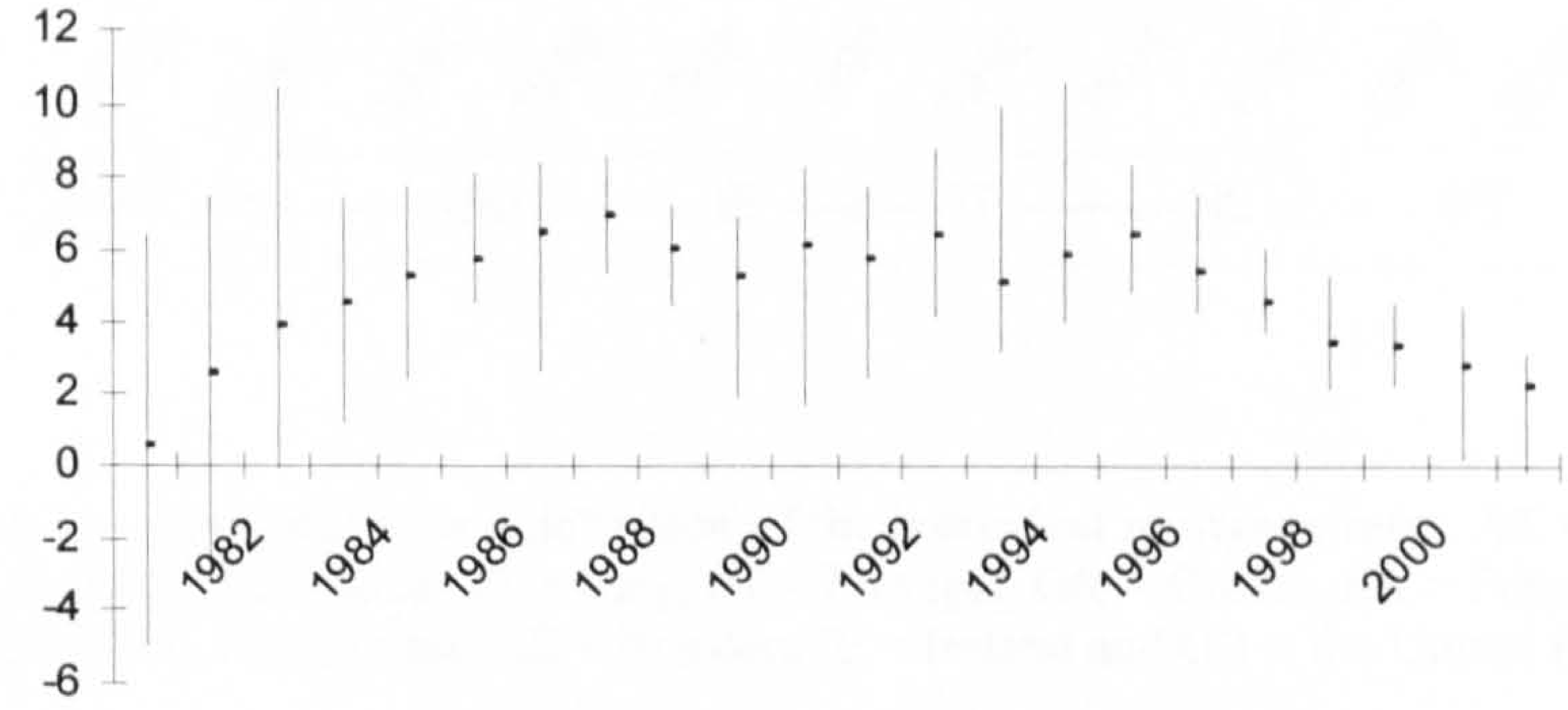
Besides nominal fluctuations, however, in order to explain long-term house price dynamics, the real cost of borrowing should be considered to be a crucial variable. Similar to nominal rates, in the 1990s real interest rates showed a gradual downward trend (Figure 3.5). While the convergence period towards the EMU has led to homogeneous nominal interest rates, however, the real rates are still quite different across EU countries due to country-specific inflation rates. In this respect, from Figure 3.5 it is easy to verify that the distance between the lowest and highest real rates at the end of the sample period is very similar to the ones witnessed at the end of the 1980s and the beginning of the 1990s. In particular, in those countries with the highest inflation rates (e.g. Ireland, the Netherlands and Spain), relatively lower real interest rates can be observed. This might partially explain the very significant property price booms witnessed in the above economies. Although the level of the real mortgage interest rates is not directly comparable across countries, they show a general pattern which is very close to the real long-term interest rates (Figure 3.8).

FIGURE 3.4 Nominal Long Term Interest Rates



Notes: Highest and lowest nominal rates. The black point is the simple average. Portugal and Greece are not included because long term rates are only available from 1985 and 1993, respectively.

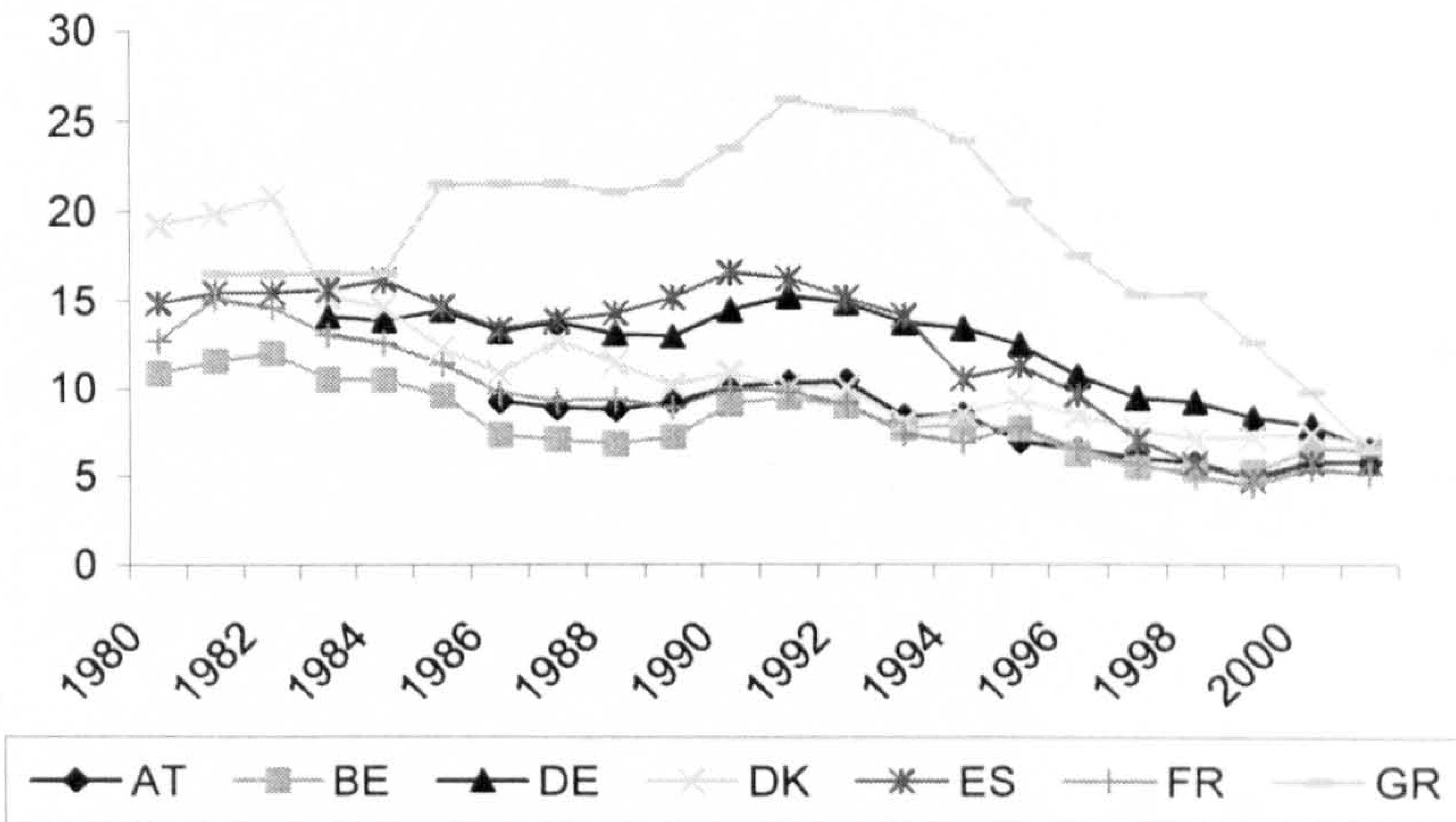
FIGURE 3.5 Real Long Term Interest Rates



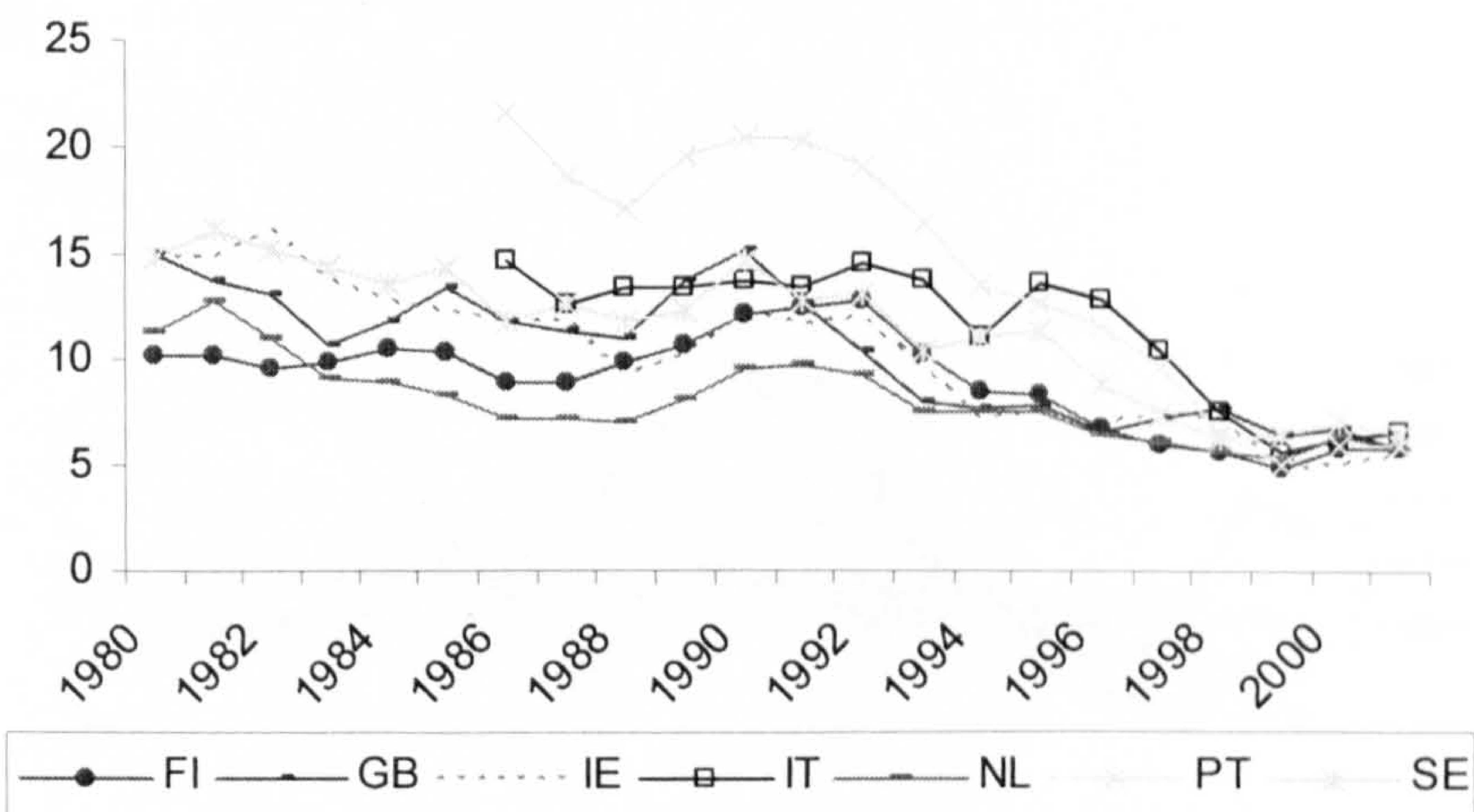
Notes: Highest and lowest real long term rates. The black point is the simple average. Portugal and Greece are not included because long term rates are only available from 1985 and 1993, respectively.

FIGURE 3.6 Nominal Mortgage Interest Rates in the EU

Panel A



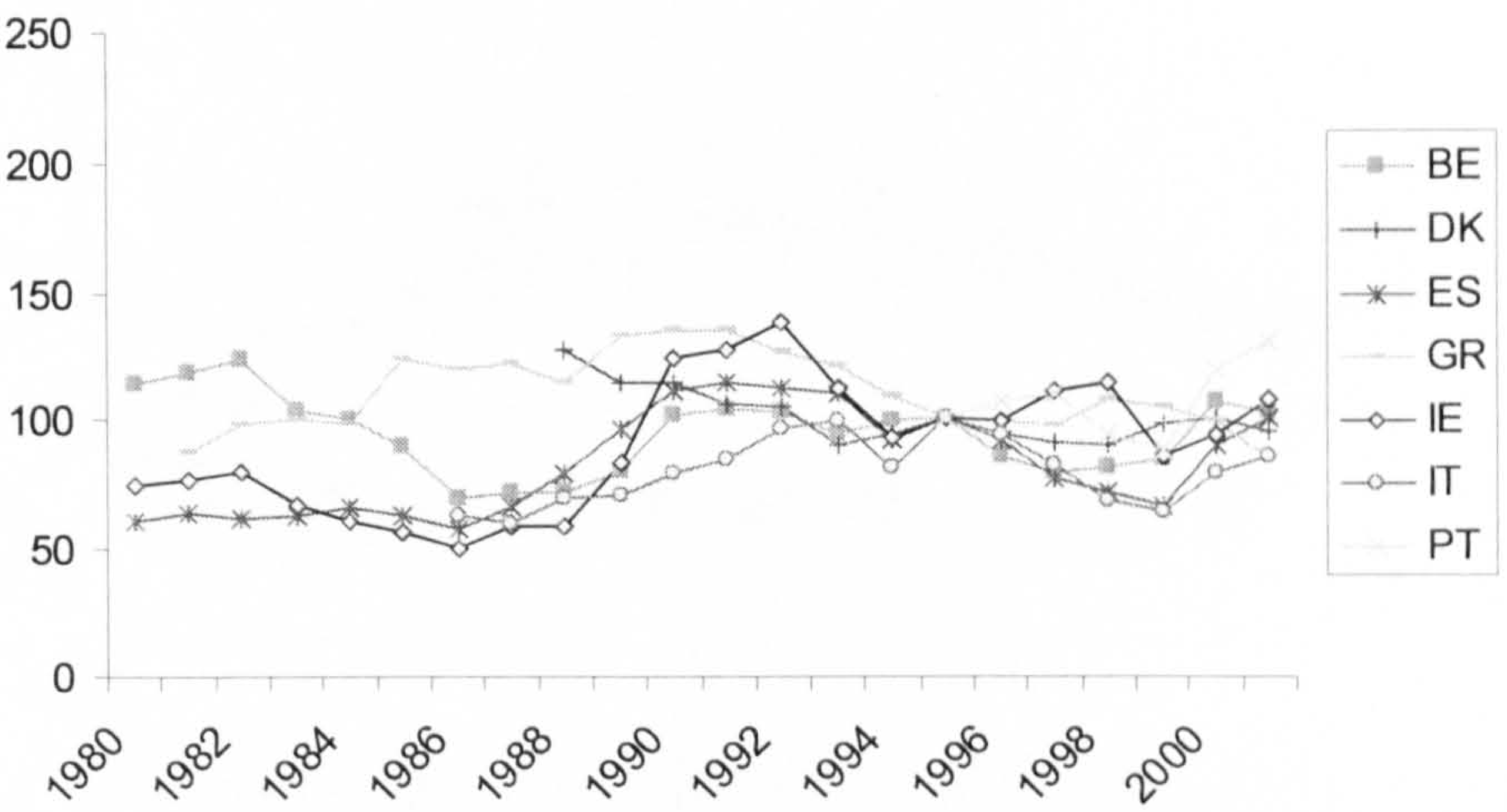
Panel B



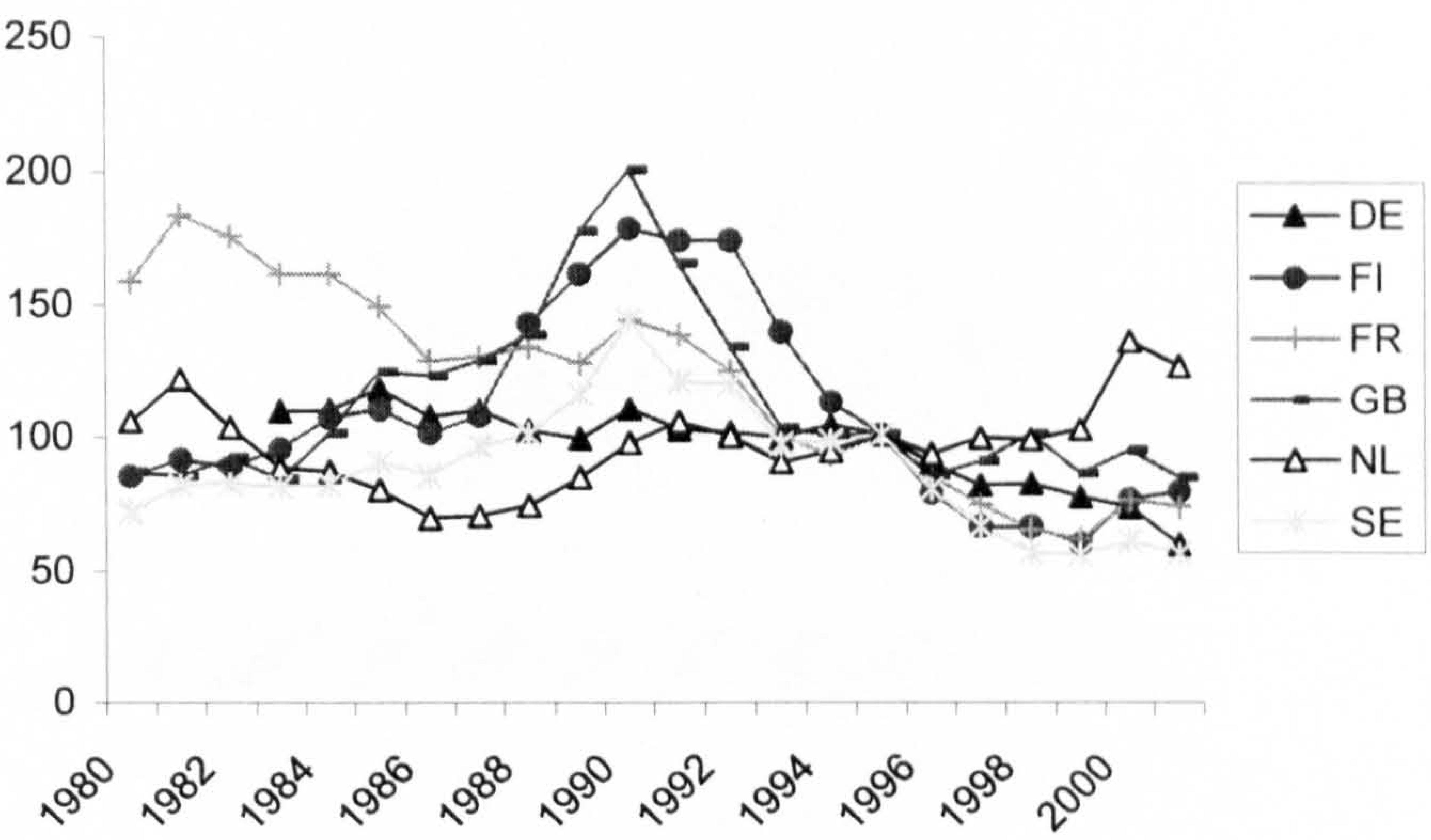
Sources: See Appendix 1 for a definition of the individual mortgage rates. AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, PT = Portugal, GR = Greece, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom.

FIGURE 3.7 Estimated Ratio of Mortgage Debt Servicing to Disposable Income (Index 1995=100)

Panel A



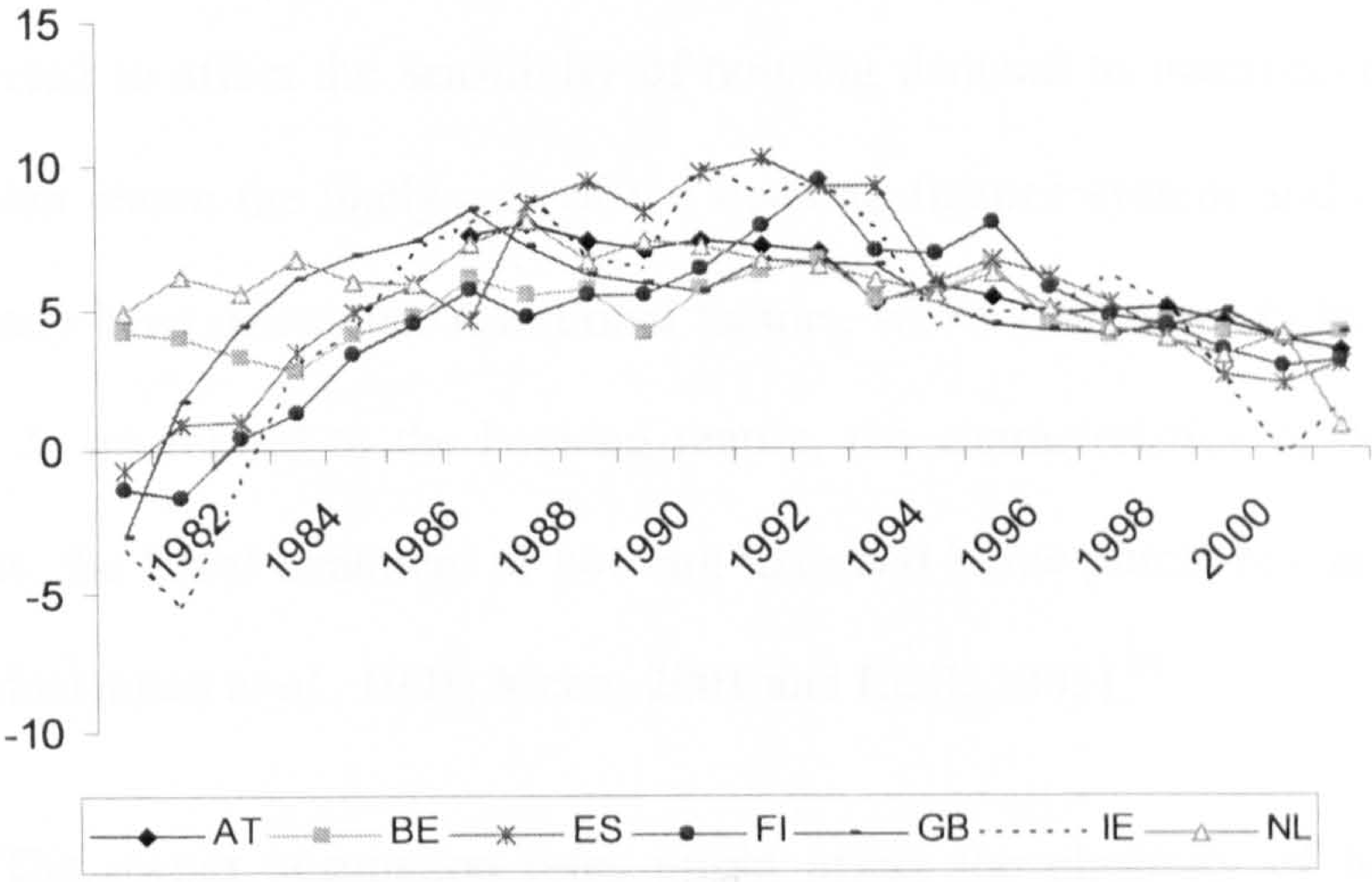
Panel B



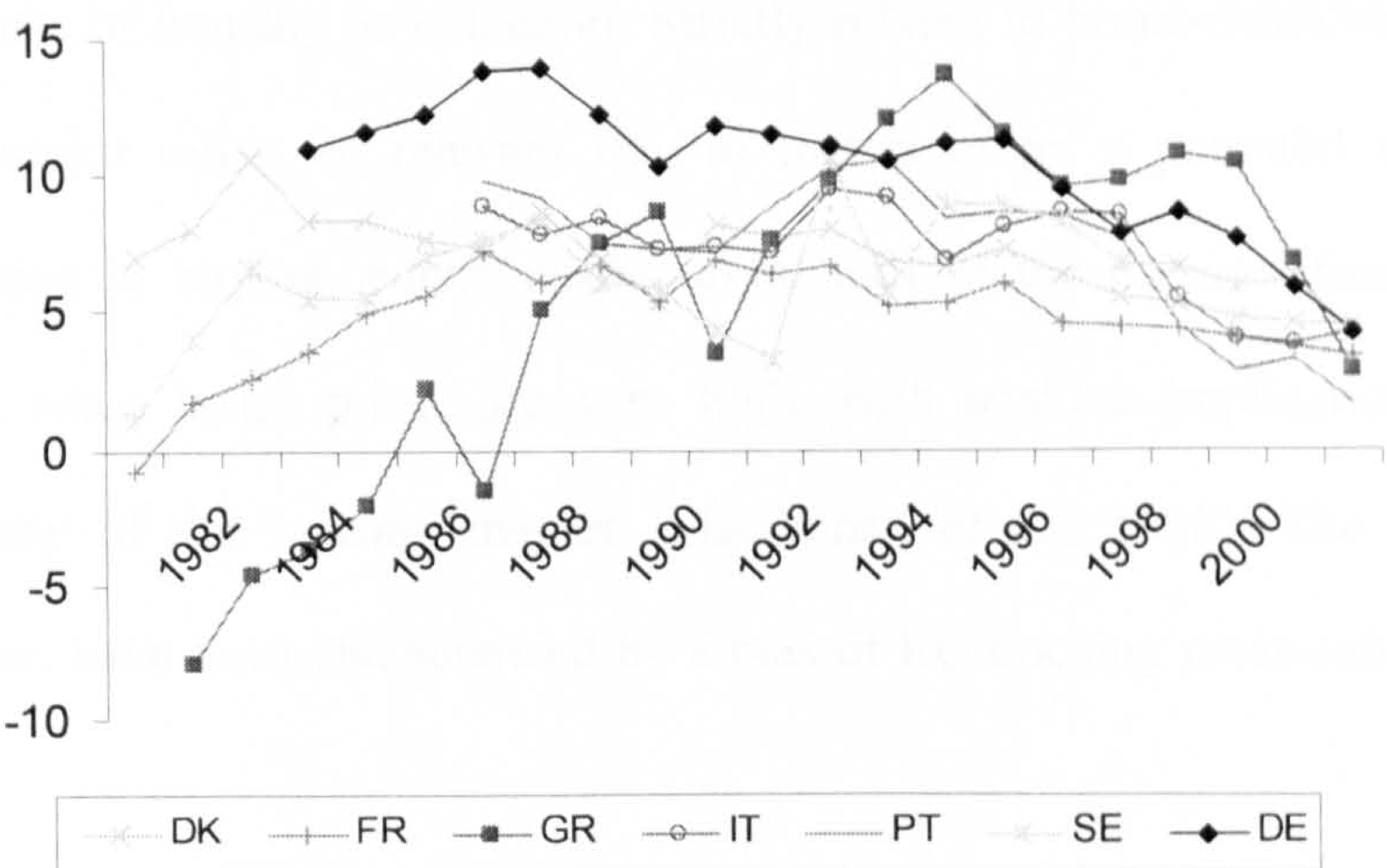
Sources: See Appendix 1 for a definition of the individual mortgage rates. AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, PT = Portugal, GR = Greece, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom.

FIGURE 3.8 Real Mortgage Interest Rates in the EU

Panel A



Panel B



Sources: See Appendix 1 for a definition of the individual mortgage rates. AT = Austria, BE = Belgium, DE = Germany, ES = Spain, IT = Italy, PT = Portugal, GR = Greece, FR = France, NL = the Netherlands, DK = Denmark, FI = Finland, SE = Sweden, IE = Ireland and GB = the United Kingdom.

2.4 *Some Additional Stylised Facts*[13]

Besides country-specific mortgage availability and interest rate developments, European housing and mortgage markets are very different in a number of other aspects which can be expected to affect the sensitivity of housing demand to macroeconomic shocks. In the section above the likely role of the housing finance system and credit availability has already been discussed. Additional factors, which have already been pointed out in chapter 2, are related to the housing tenure, the characteristics of the mortgage debt contracts, the fiscal treatment of housing loan and home purchases, and the transaction costs (Maclennan *et al.*, 1998; Meen, 2001 and ECB, 2003).[14]

The owner occupation rates might affect the elasticity of house prices with respect to interest rates indirectly, through the households' perception of wealth and borrowing possibilities. Credit constrained agents in countries with high owner occupation rates could be more sensitive to interest rate changes, in that they are aware of the use of housing as collateral. Strictly related to home-ownership, the size of the rental sector might be relevant too, in that it offers a potential safety-valve or an alternative to buying, which might divert part of the demand from owner-occupied market when house prices are very high, with positive implications for the overall efficiency of the housing market (Maclennan *et al.*, 1998). The last two decades, however, have been characterised by a bias of EU housing taxes-subsidies in favour of

[13] This section provides some additional aspects of the EU housing markets, which are relevant to our analysis. In particular, the housing taxes and the contractual features of the new mortgages are discussed.

[14] An additional point is worth making at this stage. The literature generally focuses on the elasticity of house prices with respect to mortgage interest rates or the user cost of housing. From a policy perspective, however, it is important to evaluate the pass-through from official interest rates to lending rates, which clearly might depend upon the degree of competition of the mortgage market, the legal framework and integration of financial markets (see Mojon, 2000 for a discussion). This question has been already tackled in chapter 2, with a direct estimation of the sensitivity of house prices to shocks of policy-controlled interest rates.

owner occupation (ECB, 2003). These distortions have caused a gradual reduction of the share of the rented sector over the total dwelling stock.[15]

Besides the availability of housing credit, factors affecting the sensitivity of households' demand for housing to shocks are the typical characteristics of the mortgage contracts, such as loan-to-value ratios, the diffusion of fixed versus variable mortgage interest rates, the typical maturity, and, most importantly, the fiscal treatment of interest payments, housing capital gains, stamp duties, etc.

As emphasised in the previous chapter, the financial deregulation process of the 1980s and 1990s did not occur to the same degree in all European countries. Moreover, the legal and regulatory framework related to the mortgage system differs quite significantly in Europe (Stephens, 2000). This partly explains the wide heterogeneity in the outstanding mortgage debt stock over GDP and its contractual features (that is, the mix between variable and fixed interest rates, the typical LTV ratios and the maturity) which are shown in Table 2.3 of chapter 2.

It is, however, very difficult to provide a typical contract characterising each European country. The rapid integration and competitive pressures in the house credit system, brought about big changes in the range of mortgage products available in the market, the LTV requirements and other lending criteria, as well as the length or maturity and the mix between variables and fixed interest rates. Additionally, some of these aspects are very much dependent upon the stance of the business cycle, and current and future expected interest rates.

[15] Table 3.1 in ECB (2003) shows that in most EU countries, the share of rented dwelling stock had decreased since 1980. Two main explanations have been provided. On one side, there are the strict rent controls which have reduced the supply of rented accommodations. On the other side, greater access and lower costs of mortgage credit as well as positive expectations of capital gains and tax-subsidies distortions in favour of owner occupation have been suggested.

Aware of the above considerations, Table 3.4 shows some recent figures related to the interest rate adjustment, length of the contracts and estimated average LTV ratios of new mortgage loans issued at the end of 2001 and beginning of 2002. Although data are missing for a number of countries, it is possible to verify whether there have been any convergences or significant changes in the typical contractual terms offered by credit institutions or chosen by individuals with respect to the features of the existing mortgage debt stock outlined in the previous chapter.

As far as the interest rate adjustment and the typical maturity are concerned, there have not been dramatic variations.[16] Some countries (e.g. Belgium, Sweden and Germany), however, seem to make increasing use of fixed interest rates, others (e.g. Finland, Greece, Ireland and Denmark) of variable or reviewable contracts. The maturities have remained stable or tended to increase, enhancing the debt servicing possibilities of households. A more heterogeneous picture emerges from the average LTV ratios. Despite the fact that legal and regulatory limitations on the maximum ratios are not existent or very high in most countries (around 80%), it is interesting to notice the case of the Netherlands (112%), and other countries like Spain, Sweden and Greece which show significant increases. In the period under study, on the other hand, new mortgage loans present lower average LTV ratios in the U.K., Ireland and Austria. At least for the former two countries, this may be due to the very high house prices witnessed in the last few years and the expectations of future slowdowns from lending institutions.

Table 3.5 summarises the fiscal treatment of housing loans, home purchases and other housing related expenditures which is expected to exert strong influences on the

[16] It is important to notice that in general there are no legal or regulatory restrictions on interest rate adjustment. The only exception is France where the variability cannot be less than one year.

user cost of housing services and, consequently, ownership decisions and house prices (Cecchetti and Rupert, 1996). Housing taxes across Europe differ quite significantly, not only in the mortgage interest rate deductibility, but also in one-off costs like stamp duties and running costs as taxes on imputed rents, real estate or/and VAT on repairs. Also the fiscal regime on capital gains, which may have implications for housing transaction decisions, is very diverse. Unfortunately, it is almost impossible to provide reliable measures or estimates of housing taxes which could be used for cross-country comparisons. This is made even more difficult by the fact that during the last few years housing tax policies in most EU countries experienced substantial changes, which somehow reduced the distortions in favour of immovable property (Table 3.6) through restricted tax exemption related to mortgage interest payments and introduction or gradual increase of real estate taxes (ECB, 2003). In general, however, housing policies tend to favour and promote home ownership, especially for low-income households. In particular, with the exception of Austria, Germany, France and the U.K., some form of tax deductibility of mortgage interest payments is available in all economies. EU housing policies also favour owner occupation with taxes on capital gains which are either non-existent for main residences (principal owner-occupied dwellings or 'pood') or imposed only if the house is resold within a fixed time period of 5 or 10 years. The presence of taxes on imputed rents for owner-occupiers is more heterogeneous, although the lack of reliable figures makes it difficult to quantify their relative importance across countries. Finally, although VAT on repairs and maintenance costs are present for most economies, they are generally more generous than standard ones.

Finally, a particular if not crucial role for the efficiency of the housing market is played by the burden of transaction costs, which include not only stamp duties, but also agent fees, mortgage costs and VAT on new homes. From the figures shown in the previous chapter, it emerged that a high degree of heterogeneity, which could explain the sensitivity of individual house buyers to real or financial shocks, with consequences not only on the volatility of house prices and the probability of speculative frenzies, but also on the mobility of the labour market. The last column of Table 3.5 reports one of the main components of total transaction costs, that is stamp or registration duties. They are relatively lower in the U.K., Denmark, Sweden and Finland (where first buyers are exempt). On the other side of the spectrum, transaction duties can be as high as 10% or more in Belgium, Luxembourg, and Greece.

TABLE 3.4 New Mortgage Loans Characteristics in the EU Countries

	Interest Rate Adjustment	*Usual Length of Contracts (years)*	*Legal and Regulatory Limitations on LTV*	*Estimated Average LTV Ratio*
Austria			80% or 100% if unsecured	60%(-)
Belgium	V(6%) N (19%) (-) F (75%) (+)	20 (+)	No	80-85%(=)
Finland	F(2%) V (97%) (+)		No	75-80%(=)
France			60% for the loan to be eligible to the mortgage market	
Germany	Mainly N and F (+)	Up to 30 (=)	60% LTV for loans backed by mortgage bonds	70%(=)
Greece	F (5%) V (60%) (+) N (15%)		No	70-80%(+)
Ireland	V(70%) (+) The rest mainly N		No	60-70%(-)
Italy	F (28%) (-)		80% (100% if garanteed)	
Netherlands			No	112%(+)
Portugal	Mainly V (=)		No	70-80%(=)
Spain	V (more than 75%) (=)	15-25 (+)		80%(+)
Denmark	V (15%) (+) N(10%) F (75%) (-)	30 (=)		80%(=)
Sweden	F(38%) (+) V(24%) N(24%)			80-90%(+)
United Kingdom	V (72%) (=) N (28%) (=)			70%(-)

Sources: ECB (2003), Table 5.1, pp. 50. Figures refer to new mortgage loans at the end of 2001 and beginning of 2002. Fixed (F): rate fixed for more than 5 years or until maturity; Renegotiable (N): rate not fixed over entire term, but more than one year and up to 5 years; Variable (V): after 1 year interest rate renegotiable or tied to market rates or adjustable at discretion of lender. (-), (=) and (+) indicate a decrease, the same and an increase in the underlying aspect with respect to the features of the existing mortgage debt stock displayed in Table 2.2 of Chapter 2.

TABLE 3.5 Housing Taxes in EU Countries in 2002

	Tax Deductibility of Mortgage Payments	*Tax on Capital Gains*	*Tax on Imputed Rent*	*VAT on Repairs (and New Homes)*	*Stamp Duty*
Austria	No in 1998	Yes (turnover <10 years)	No	10-20% (10-20%)	6%
Belgium	Yes	Yes (turnover <5 years, exemption for owner occupiers)	Yes	6%, 21% (21%)	10%-12.5% registration fee
Finland	Yes	Yes (exemption for pood after 2 years)	Yes	22% (22%)	4% of purchase values (exemption for first-buyers)
France	No	Yes (exemption for main residence)	No	5.5% (19.6%)	2-3% (7-10% in 1994)
Germany	No	Yes (turnover <10 years, exemption for owner occupiers)	No	16% (16%)	3.5%
Greece	Yes (for pood)	No	Yes (for pood)	18% (0%)	11-13%
Ireland	Yes	Yes (exemption for main residence)	No	12.5% (12.5%)	0% - 9%
Italy	Yes (only for pood)	Yes (50% tax deduction for pood). Abolished 2002	Yes (exception for pood)	10% (4%) for pood 19% (19%) for others	4% (pood) Up to 9% (others)
Luxembourg	Yes	Yes (exemption for pood)		3% (3%)	7-10%
Netherlands	Yes	No	Yes	19% (19%)	6%
Portugal	Yes	Yes (exemption if proceeds are reinvested in another residence within 2 years)	No	5%-17% (0%; municipal transfer tax 0% - 10%, progressive)	10% in 1994 0.80%
Spain	Yes	Yes (exemption for principal dwellings when reinvested)	No (for primary houses)	15% (6%; 7%)	6% in 1994
Denmark	Yes	Yes (exemption for owner occupiers)	Yes	25% (25%)	1.5%
Sweden	Yes	Yes (25%)	Yes	25% (25%)	1.5% 3%
United Kingdom	No (abolished in 2000)	Yes (tax exemption for pood)	No	17.5% (0%)	1%, 2%, 4% depending on house value

Sources: Maclennan, *et al.* (1998), Henley and Monley (2001) and ECB (2003).

Notes: Data refers to 2002 unless differently specified. 'Pood' means principal owner-occupied dwelling.

TABLE 3.6 Major Reforms of Housing Tax and Subsidy Policies in the 1990s

Belgium	Restriction of property tax reductions to owner-occupiers; continuous increase of transaction taxes (VAT) on new buildings
Denmark	Removed tax incentives for housing investment; continuous increase of stamp duties
Germany	Several measures to reduce tax deductions and subsidies for investment in housing; abolished property tax; increased transaction costs
Greece	Raised property tax and transaction costs through increased administrative value of real property for tax purposes. Introduced new tax on large real estate property
Spain	Reduced tax deductions for secondary and rented dwellings; abolished imputed income on principal dwellings
France	Raised property tax; abolished taxes on imputed rents; tax reduction for low-income households; reduced transaction taxes
Ireland	Reduced tax deductions for interest payments, abolished for landlords; abolished property tax and halved capital gains tax
Italy	Introduced local property tax; tax reductions for owner-occupiers and some categories of landlords; reduced registration tax for owner-occupiers
Netherlands	Reduced tax reliefs for interest payments and restricted it to principal dwelling; introduced subsidies for low-income first-time buyers
Austria	Reduced indirect subsidies; set-up of housing construction banks
Portugal	Restriction on and latter the end of mortgage subsidies for new loans
Finland	Introduced state-guarantees for owner-occupiers loans
Sweden	Increase of property tax rate; tax reform to neutralise incentives for different forms of housing investment
United Kingdom	Phased out interest relief system

Sources: ECB (2003), Table A.2, pp 53.

3. *Modelling National House Prices*

3.1 *The Demand-Supply and the User-Cost Specifications*

The housing economics literature modelling of house price dynamics is based on two approaches, both leading to very similar specifications.[17] In the first approach, an equilibrium house price equation is derived by equating housing demand and supply functions. In particular, following the early paper by Nellis and Longbottom (1981), a general demand for housing could be given by the following implicit specification:

(1) $$H_t^d = D(g_t, y_t, dem_t, mr_t, ms_t)$$

and a supply function by:

(2) $$H_t^s = S(g_t, h_{t-1})$$

where g_t is the real house price, y_t the real income, dem_t is a demographic variable, mr_t is the mortgage interest rate , ms_t is the mortgage stock and h_{t-1} is the previous period's housing stock.

In equilibrium $H_t^d = H_t^s$, which leads to the following house price equation:

(3) $$g_t = f(h_{t-1}, y_t, dem_t, mr_t, ms_t)$$

where $\frac{\partial g}{\partial h}, \frac{\partial g}{\partial mr} \leq 0$ and $\frac{\partial g}{\partial y}, \frac{\partial g}{\partial dem}, \frac{\partial g}{\partial ms} \geq 0$

Eq. (3) represents the conventional reduced-form house price specification estimated by the early (Hendry, 1984 and Nellis and Longbotton, 1981) as well as the recent

[17] Meen (2001) points out four classes of house price models identified in the literature: early *ad hoc* models with very limited theoretical structure, 'mark-up' models relating house prices to construction costs, reduced form models and life-cycle models. On the basis of the recent empirical studies and the similarity of the proposed house price equations, the latter two approaches are reviewed.

(Kasparova and White, 2001 and Barot, 2001) housing market literature. The expected signs of the partial derivatives have a straightforward interpretation.

The second approach, which is the starting point for many house price models (see Breedon and Joyce, 1993; Muellbauer and Murphy, 1997; Pain and Westaway, 1997; Meen, 1998, 2000, 2002 amongst others) is based on the life-cycle model of Buckley and Ermish (1982) and Poterba (1984). This approach captures both the consumption and investment features of housing demand through a formal derivation of the user cost of housing.

In particular, a representative household could be assumed to maximise the expected value of a time-separable, time-invariant lifetime utility function depending upon the number of housing units H_t owned at the beginning of time t and the composite consumption good c_t:[18]

$$(4)\qquad Max\ E_0(U_0)=\sum_{t=0}^{T}\left[\left(\frac{1}{1+\rho}\right)^t U\left(H_t,c_t\right)\right]$$

subject to the intertemporal budget constraint:

$$(5)\qquad \begin{aligned} c_t = y_t + (1+r)S_{t-1} - S_t - (1+r_m)M_{t-1} + M_t \\ - g_t\left(H_t-(1-\delta)H_{t-1}\right) - g_tH_t(\tau_p+\mu) \end{aligned}$$

which can also be written as:

$$(5.1)\qquad \begin{aligned} c_t + S_t + (1+r_m)M_{t-1} + g_tH_t(1+\tau_p+\mu) = \\ y_t + (1+r)S_{t-1} + M_t + g_t(1-\delta)H_{t-1}\ ,\forall t \end{aligned}$$

[18] Miles (1994) argues that, in constructing theoretical models for analysing the housing market, one is forced to make some strong assumptions which abstract from real characteristics of the housing market. In particular, the three main assumptions featured in this models are: individual houses can be compared and aggregated (i.e. a housing unit is of standardised size and quality); preferences over houses and other commodities are constant, and household labour incomes are independent of housing, implying that wages and supply of labour are unaffected by house value changes and type of housing. While it is clear that these factors might be relevant, Miles argues that the main qualitative predictions of the model are unchanged.

The left-hand side of (5.1) provides the uses of funds, with respectively the current consumption (c_t), the net financial assets (*excluding* mortgage debt) at the beginning of time t (S_t), the real expenditures on mortgage principal and interest (where r_m is the real interest rate in terms of consumer goods paid on mortgages) and the current purchase of housing stock ($g_t H_t$ with g_t being the price, in terms of consumer goods, of a unit of housing at time t) augmented with expenses relative to home ownership (where τ_p is the property tax rate and μ the constant fraction of home value necessary for insurance, maintenance and repairs costs). The right-hand side defines the sources of funds: real labour income in terms of consumer goods (y_t), the current stock of net financial assets gross of real interest rate earning (r), the outstanding mortgage at the beginning of time t (M_t) and the current value of housing stock net of depreciation (δ).

According to this maximisation problem, agents are assumed to earn income from their fixed supply of labour and receive (or pay) interest on net financial assets. Agents, however, are assumed to be able to borrow against housing wealth. This assumption is central, in that the mortgage borrowing constraint can be imposed:

(6) $M_t = \beta g_t H_t$

where β is the loan-to-value ratio which lies between zero and one. Condition (6) implies that mortgage borrowing fluctuates with house prices, *ceteris paribus*. The implication of the equality constraint is that consumers are forced to pay off principle whenever house price falls, an assumption which is not too realistic.

The model can easily be solved by substituting (6) into (5) and maximising (4) subject to (5) with respect to c_t, S_t and H_t. The corresponding first order conditions (FOCs) yield the following relationship:

$$(7) \qquad \frac{U_H'(H_t,c_t)}{U_c'(H_t,c_t)} = g_t\left[(1-\beta)+(\mu+\tau_p)+\left(\frac{1+r_m}{1+r}\right)\beta-(1+\dot{g}_t^e)\left(\frac{1-\delta}{1+r}\right)\right]$$

where $\dot{g}_t^e = \frac{g_{t+1}-g_t}{g_t}$. Equation (7) provides the condition, according to which, at optimum, the marginal rate of substitution between housing services and consumption goods equals the *user cost* of housing and represents the slope of the budget line (i.e. the amount of current consumption to sacrifice to purchase one unit of housing). In the special case, where $r_m = r$, (7) simplifies to:

$$(8) \qquad \frac{U_H'(H_t,c_t)}{U_c'(H_t,c_t)} = g_t\left[(1+\mu+\tau_p)-(1+\dot{g}_t^e)\left(\frac{1-\delta}{1+r}\right)\right],$$

which (for small values of δ, r and $\dot{g}_t^e$) can be approximated by:

$$(9) \qquad \frac{U_H'(H_t,c_t)}{U_c'(H_t,c_t)} = g_t\left[r-\dot{g}_t^e+(\tau_p+\mu+\delta)\right]$$

Eq. (9) is the familiar formula in which the right hand side is the real user cost of housing capital (ruc_t). As comprehensively pointed out in Meen (2001), the above definition can be extended in a number of ways. It might include the household marginal income tax, transaction costs, and terms allowing for the impact of credit constraints.[19] Most of the empirical literature uses different specifications according to

[19] In his seminal paper, Ermisch (1984) introduced credit constraints in the user cost of housing deriving the following condition $\frac{U_H'(H_t,c_t)}{U_c'(H_t,c_t)} = g_t\left[r-\dot{g}_t^e+(\tau_p+\mu+\delta)+\lambda_t/U_c'(H_t,c_t)\right]$ where λ_t represents the shadow price of rationing constraints.

the specific fiscal treatment of mortgage interest rates deductibility, as well as data availability.

A general formulation for the real user cost of housing services, which includes tax deductibility of nominal interest rate payments, gearing and the opportunity cost of internal financing is:

(10) $ruc_t = g_t \left\{ \beta_t (1-\tau_t) i_{mt} - (1-\beta_t)(1-\tau_{it}) i_{ft} - (\pi_t^e + \dot{g}_t^e) + (\tau_{pt} + \mu_t + \delta) \right\}$

where β is the proportion of housing expenditure financed through mortgage debt, whereas $(1-\beta)$ is the part financed with internal resources, τ is the effective tax relief on mortgage interest payments, τ_i is the income tax rate, i_m is the nominal mortgage rate, i_f is the nominal rate of return on invested funds.[20] The variable π^e is the expected rate of inflation and $\dot{g}^e$ the expected real house price appreciation. The term $(\pi^e + \dot{g}^e)$ can be thought as the nominal expected house price capital gains.

The above definition of real user cost of housing services (which treats housing both as a consumption and an investment good) would be the ideal measure to include in any empirical study to modelling house price dynamics. However, with the only exception of very few studies mainly focusing on the United Kingdom, there are not available (and reliable) data on marginal tax rates and LTV ratios, costs of maintenance, and depreciation rate of housing structure. While the latter components could be assumed to change only gradually over time, representative housing and income tax rates, and loan-to-value ratio for home buyers are not available for most countries. As a result, only a simple approximation of the user cost is considered below. In particular,

[20] Some authors include a transaction cost component, which is defined as the sum of transaction cost including estate agents fees, legal costs and stamp duty. This is then divided by the average holding period of a house and scaled up to allow for discounting (Breedon and Joyce, 1993).

the two main components are used, namely the real interest rates (r_t) and the real housing capital gains ($\dot{g}_t^e$).

Using a simplified formulation of the user cost given in eq. (9) (that is, ignoring the terms τ_p, μ and δ) with time dependent real interest rates, housing market efficiency requires also the following arbitrage condition to hold:

(11) $R_t = g_t\{r_t - \dot{g}_t^e\}$ or $g_t = R_t / \{r_t - \dot{g}_t^e\}$

where R_t is the real imputed rental price of housing services, i.e. the amount of money to be given to a household to compensate of the loss of one unit of housing. The imputed rental price of housing could be proxied with the market rent for a property of similar quality to the average owner-occupied home (Miles, 1994). However, there are two main problems strictly related to the market price of the housing services. Firstly, rented properties have different characteristics. For instance, they tend to be smaller and of reduced quality than the average owner-occupied home. Secondly, in most countries the rented sector has been or is highly regulated, so that the recorded private rents are not free market prices.[21]

For empirical purposes, therefore, R_t is unobservable and, due to data deficiencies, substituted out in terms of its determinants in the literature. Most authors (among which, Meen, 1990, 1998, 2000, 2002 and Breedon and Joyce, 1993 argue that the real rental price can be assumed to be a function of the demand for and supply of housing services. At an individual level, the demand for housing services might depend on real income (y_t) and financial wealth (W_t), whereas, at aggregate level, the

[21] Although the evolution of rental market has strong implications for the functioning of the housing market, this chapter does not attempt to disentangle the role of rents for house price dynamics. See ECB (2003) for a clear discussion.

exogenous rate of household formation (dem_t) might constitute an additional factor. The supply of housing services can be thought of as being proportional to the existing stock of dwellings H_t.

As a result, the market clearing rental price is generally defined by:

(12) $R_t = f\left(y_t, W_t, H_t, dem_t\right)$

Substituting eq. (11) into (12) and assuming a logarithmic specification for R_t yields the following general formulation:

(13) $\ln(g_t) = f\left[\ln(y_t), \ln(dem_t), \ln(W)_t, \ln(H_t), \ln(ruc_t)\right]$

The fact that the structural relationship given by equation (11) is lost in eq. (13) depends upon the unknown factors driving the rental price of housing in eq. (12) and the specific availability of relevant data.[22] As a result, the house price specifications proposed in the housing literature are very different from each other. For instance, lack of data on loan-to-value ratios, marginal income tax rates, property taxes, depreciation rates and maintenance costs constrains many researchers to focus on simplified versions of the real user cost of capital. Moreover, the latter might take negative values, because high inflation and housing capital gains expectations can outweigh the costs of housing. This implies that it is impossible to log-transform the series as eq. (13) would suggest. Although correcting for a positive and fixed risk premium on housing and the impact of

[22] The house price equations specified in equation (13) and equation (3) are very similar to each other, in that they both account for income, demographic and housing stock dynamics. They differ, however, in a number of aspects. The reduce-form specification includes a measure of housing lending stock which is not present in the user-cost equation, whereas the latter incorporates a measure of household wealth which is assumed to affect the permanent income. More importantly, eq. (3) does not include a full measure of the user cost of owing a house, which is mainly driven by the cost of borrowing (that is the mortgage interest rate) and the housing capital gains/losses.

credit restrictions might solve the problem (Miles, 1995 and Meen, 2002), most empirical studies include the user cost in levels for estimation purposes.[23]

As pointed out in Meen (2001) and Meen and Andrew (1998a), no empirical paper estimates the static long-run relationship provided in (3) and (13). This is due to the fact that search and transaction costs are very substantial in the housing market. As a result, the most typical approach used in the British and North European housing literature is based on dynamic specifications, which is the method used in this chapter. Moreover, given that the main difference between the demand-supply and the life-cycle-motivated models turns out to be based upon the inclusion of the expected housing capital gains in the user cost (which in all but very few exceptions are approximated assuming extrapolative expectations),[24] it is evident that distinguishing empirically between the two approaches is very difficult.[25] On the basis of these considerations, the empirical section showing the individual country specific equations will provide the results of two separate house price equations: the first with the inclusion of a real user cost proxy (given by the difference between real mortgage rates and real house capital gains), the second with the real mortgage rate only. Before that, however, a brief summary of the existing cross-country empirical literature with some of its limitations is outlined.

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[23] Additionally, to take into account the role of tilting effects, some studies split real user cost into nominal user cost (i.e. the after tax nominal mortgage rate) and expected nominal gains on housing.
[24] This assumption is further discussed below.
[25] At this stage it might be interesting to compare the specifications given by the demand-supply and life-cycle approaches with the ad-hoc and mark-up models mentioned above. The former characterised the U.K. housing literature in the seventies, during which strong booms in house prices were thought to be the effect of strong growth in mortgage credit. In particular, beside standard determinants as income and mortgage interest rates, their main feature was the inclusion of net or gross mortgage advances or stock of credit to control for credit market restrictions. While this was valid as long as the market was rationed, since with the financial liberalisation credit became endogenous to the model, this approach is not too well suited for modern house price equations. This also applies to reduced-form models which include the contemporaneous measures of mortgage stock. The mark-up models, on the other hand, were based on the assumption that if the supply of housing is perfectly elastic then in the long run house prices should be formed as a mark-up on construction costs. The latter assumption, however, is not supported by the data and this approach never took off in the European literature.

4. A Review of the Existing Cross-Country Studies of House Prices

While the empirical literature focusing on national house price dynamics is very extensive,[26] relatively little research is available at a cross-country level. The only exceptions are Kennedy and Andersen (1994), Englund and Ioannides (1997) and, more recently, Kasparova and White (2001). This section reviews the latter studies from which the empirical analysis carried out below builds on.

Kennedy and Andersen (1994) model the booms and busts in house prices by analyzing movements in the ratio of house prices to income in 15 OECD countries over the period 1970-1992. Similarly to a number of previous studies carried out at national level, they specify and estimate a dynamic house price equation, in which the explanatory variables are: the level of real household disposable income; the unemployment rate – to capture distributional effects as well as being a potential proxy for income uncertainty; the user cost of housing, constructed as the real mortgage interest rate less the real housing capital gains (measured by the lagged real house price inflation rate); the ratio of the 15 to 64-year-olds as a proportion of the total population to control for demographic changes; the lagged ratio of household debt to income to proxy for financial liberalisation; a time trend to proxy the housing stock and, finally, the lagged dependent variable. Using annual data, they estimate this equation separately for each of the OECD-15 and provide support for the presence of different determinants of house price dynamics across countries. In particular, the user cost proxy is found to be negatively signed and statistically significant in eight countries. In most of the others, the real interest rate performs better. While the lagged dependent variable is significant in all regressions but the U.K., the demographic term has a predictive power only in

[26] See Meen (2001) for a recent survey on national house price studies carried out in the U.K., the U.S. and Scandinavian countries.

France and the United States. Finally, the financial liberalisation variable is significant in Japan, the U.K. and Norway. The time trend is found to be negative in four countries with highly significant coefficients in Norway and the United Kingdom. This is consistent with the latter picking housing stock dynamics.[27]

Englund and Ioannides (1997) (thereafter EI) compare the dynamics of house prices for the same panel of countries over the same sample period. By using panel data estimation techniques, they specify a dynamic regression, where the dependent variable is the real house price change and the three explanatory variables are lagged house prices, the rate of growth of real GDP and the rate of change of real interest rates (ex-post and pre-tax). They also test for the role of the rate of change of population in the house-buying ages (i.e. 20-30). The latter variable is not statistically significant and not included in the reported specifications. They first examine whether one-period-lagged changes of income and interest rates have a predictive power in explaining current house price changes. They find that in addition to lagged price changes, both GDP and real interest rates have a strong explanatory power. Although the fit obtained with pooled data proves to be quite good, they test for the equality restrictions of the parameters and reject them at conventional levels of significance. As a result, they compare the pooled regression results with estimates conducted separately for each of the fifteen OECD countries. Similarly to Kennedy and Andersen (1997), they show quite remarkable cross-country differences, but, due to degrees of freedom problems, their estimates are still not statistically well determined. In particular, the coefficient associated with the GDP growth is positive and statistically significant only in Finland, Ireland, Norway and the United Kingdom, whereas lagged changes of the dependent variable are statistically significant in most of the countries.

[27] See section 5.2 for a discussion of the role of a time trend in house price equations.

In the second part of their paper, they assess what types of shocks are important in explaining current house price movements including current changes of the explanatory variables. Similarly to the previous model, GDP growth turns out to be very significant, varying between 1.15 and 1.22 according to specifications with different deterministic terms. The coefficient associated with the real interest rate has the expected negative sign and ranges between –0.016 and –0.011. Not surprisingly, their results are less satisfactory at individual country level. In the separate regressions, GDP coefficients are positive (except for Italy) and significant in 11 countries, whereas the parameter associated to the real interest rate is negative and weakly significant only in six out of fifteen. But, as the same authors argue, their results (as well as the ones produced in Kennedy and Andersen, 1994) may be biased by the presence of few degrees of freedom. An additional limitation is that they do not account for the univariate properties of the variables used in the regressions, nor for the possibility of long-run cointegrating relationships among them.

A step forward in this direction was recently taken by Kasparova and White (2001), who study the residential prices in four EU countries (Sweden, the Netherlands, the U.K. and Germany) using annual data over the period 1970-98. Similarly to Englund and Ioannides' study, they use panel data techniques, but, in contrast to them, estimate a two-step error correction mechanism (ECM) where real house prices are dependent on real GDP, housing starts (that is the number of new dwellings) and nominal mortgage interest rates. Results provide statistical significance of the first two variables (with long-run coefficients equal to 0.34 and –0.42, respectively), whereas the interest rate is found to have no effect. After testing and rejecting the equality of the GDP and housing starts coefficients, they apply the same model to each country separately. In particular, they find that the mortgage rate becomes significant in Germany, the U.K. and the Netherlands. In the latter country, however, the sign of the corresponding coefficient is positive. The long-run elasticity of house price with respect to the real GDP is significant in Germany, Sweden and the U.K. and varies between 1.45 and 0.31. They conclude by arguing that the results show a very high degree of heterogeneity and it is not appropriate to pool all four countries. In this respect, it is suggested that to account for country-specific effects of macroeconomic shocks on residential prices across countries, individual house price equations should be estimated. The presence of few degrees of freedom, however, represents a major weakness of their results. Moreover, they do not account for the financial deregulation process which strongly affected the availability of housing lending in many European countries.

4.1 Modelling Financial Liberalisation in House Price Equations

As discussed in the previous sections, many countries that experienced house price booms also went through a significant financial liberalisation process. While most authors agree that any house price equation should take these effects into account, the empirical literature has modelled the increasing relaxation of credit availability and competitiveness in the housing lending sector in different ways.

Fernandez-Corugedo and Price (2002) review the main theoretical and measurement issues related to financial liberalisation in the consumers' expenditure studies in the United Kingdom. In particular, they point to several liberalisation proxies which have been used in the literature. They are the ratio of the stock of personal sector consumer credit to GDP, the interest differential between deposit and borrowing rates, the loan-to-value and loan-to-income ratios, mortgage equity withdrawal, and the proxy FLIB suggested by Muellbauer and Murphy (1993). The latter is defined as the sum of the constant and the residuals in a regression of the loan-to-value ratio on the house price to income ratio, the nominal post-tax mortgage rate and a two-year moving average of the post-tax mortgage rate. FLIB has been widely used to account for the proportion of liquidity-constrained individuals in estimating forward-looking consumption functions (Muellbauer and Murphy, 1993; Darby and Ireland, 1994 and Caporale and Williams, 2001) and house price equations (Muellbauer and Murphy, 1997). Fernandez-Corugedo and Price (2002), however, show that FLIB is unable to accurately depict the consequences of the U.K. financial deregulation in the 1990s, in that, according to this measure, all the liberalisation that occurred in the 1980s was reversed the following decade, which is not very likely.

In this respect, they argue that, although *a priori* it is not obvious which proxy would be the most appropriate to account for the effects of deregulation, a measure of liberalisation should be expected to increase rapidly in the 1980s and just slow down somewhat thereafter. These dynamics could be captured by alternative proxies related to consumer credit and loan-to-value ratios, but they are subject to the main criticism of being endogenous and affected by many other factors other than liberalisation.

Financial deregulation strongly affected the housing finance market and it seems reasonable, if not necessary, that modelling house prices should at least make an attempt to incorporate and account for the huge mortgage credit availability changes occurred during last two decades. As shown above, some of the international studies (Englund and Ioannides, 1997 and Kasparova and White, 2001) ignore these effects completely. Kennedy and Andersen (1994), on the other hand, use the lagged changes (and levels) of the ratio of household debt to income as an explanatory variable, finding significant positive effects in some countries. The lag of households' indebtedness over disposable income is also used in Kostela, *et al.* (1992) in their Finish housing market paper and in Hendry (1984) for the United Kingdom.

While there may be alternative measures which could better isolate the pure effect of liberalisation, and disentangle mortgage loan supply from demand effects, the above broad proxies might be useful in modelling or, at least, controlling for significant innovations in the housing lending sector which occurred during the period under investigation. As a result, also due to data availability constraints, this chapter follows Hendry (1984), Kostela, *et al.* (1992) and Kennedy and Andersen (1994) and make use of two broad financial liberalisation proxies: lagged changes of the real mortgage stock (RMSTOCK) or, in order to control for real activity dynamics, the ratio of the mortgage stock to GDP (MSTOCKY). The results will be discussed in the following sections.

5. *Pooled versus Heterogeneous House Price Equations*

This section extends the previous literature based on pooled regressions comparing the dynamics of housing prices in thirteen European countries over the period 1970-2001, but formally tests for the equality assumption of the coefficients associated with the main driving factors.[28] Distinct from the individual country estimation provided in section 6, and to enhance comparability with other cross-country studies, the annual house price indexes collected by the Bank for International Settlements (BIS) are used. Appendix 1 contains a detailed summary of the definitions, which, with the exception of Germany and France, coincide with the ECB database used in the next section.[29]

The use of annual series has a number of advantages as well as drawbacks, some of which have been discussed above. First, low frequency house price indexes are available for longer periods than the quarterly or semi-annual data for most of the countries. This allows us to improve the modelling and identification of long-run relationships. At the same time, however, the use of annual time-series reduces the number of observations and constraints to estimate relatively more parsimonious specifications. Moreover, low frequency data do not permit modelling or capturing within-single-year effects which in practice might be relevant. The latter two points will be overcome in the following section, in which separate house price equations for each country are estimated.

In this section two groups of house price specifications are estimated. The first one extends the work of EI and specifies parsimonious equations of house price changes

[28] Greece and Luxembourg are not included because of data availability. For Portugal and Austria, house price indexes start only in 1988.
[29] While for France the annual BIS index refers to the entire nation, the quarterly series are available for the area of Paris only. The two series, however, are remarkably similar. For Germany the official BIS and ECB house price indexes differ significantly from each other.

on changes in real household disposable income, the change of the real and nominal interest rates and the rate of change of a demographic variable. In the second group, the non-stationarity properties of the variables are accounted for by trying to identify the long and the short-run determinants in an unrestricted error correction mechanism (ECM) framework implemented through the Seemingly Unrelated Regression (SUR) method, and the recent Pooled Mean Group (PMG) approach suggested by Pesaran *et al.* (1999a).

5.1 A Simple Extension of the EI House Price Equations

Given that the number of countries ($N=13$) is smaller than the number of observations ($T=32$), a natural estimator is provided by the SUR (or Zellner's) method. This approach allows estimation of the parameters of the system accounting for the heteroskedasticity, and contemporaneous correlation in the errors across equations producing more efficient estimates. Moreover, this method is easily applicable to more general forms of SUR, in which the parameters associated with particular explanatory variables are country specific, and allows testing for cross-equation restrictions and equality assumptions through traditional Wald tests.

This is important for a number of reasons. Firstly, as argued in Hague *et al.* (2000), neglecting heterogeneity and ignoring differences across countries can lead to biased estimates and overestimate the influence of certain factors. Secondly, European housing and mortgage markets seem to be very diverse and it might be informative to formally test for this degree of heterogeneity.

The models estimated and shown in Table 3.7 closely follow the approach of EI, who test whether, besides lagged changes of house prices, other variables have some predictive power in explaining 'future' house prices. In particular, the real house price change (DRHP) is regressed on the 'lagged' change of the real household disposable income (DRDI), the real long-term interest rate (DRLT) and the ratio of the population between 25 and 44 over the total to account for demographic factors (DDEM). All variables are in logs, with the exception of interest rates, which are in levels.[30]

Column 1 shows the results of the first specification, in which not only common slopes, but also a common constant are imposed. Similarly to EI, it is found that in addition to lagged price changes, lagged real disposable income and real interest rates changes have a strong predictive power. The corresponding coefficients are statistically significant at five percent level, and have the expected signs. The DW statistic, however, indicates the presence of some first-order serial correlation in the residuals. This specification is then augmented with an additional lag of the dependent variable (column 2). The parameter associated with the latter term is not only statistically significant at one percent level, but, as other studies have shown (Case and Shiller, 1988; Englund and Ioannides, 1997) is negatively signed. Differently from the previous model, the demographic variable becomes significant (although only weakly) and has a positive explanatory power.

To capture 'tilting' effects, the specification of column 3 is augmented with lagged changes of the nominal long-term interest rate (DLT), which are found to have strong negative effects. While the demographic variable becomes insignificant, the adjusted-R^2 rises from 0.35 to 0.38. According to the estimated parameters, one

[30] There are clearly alternative and additional variables that could enter a reduced-form house price equation of this type. For instance, mortgage interest rates could be used in place of long-term interest rates. At the same time, given the importance of mortgages for housing demand, past values of residential lending changes could be included. The latter two variables, however, are only available from the early 1980s, and in using them the sample would be reduced to less then twenty observations.

percentage faster real disposable income this year causes a 0.30% faster house price growth tomorrow, whereas a 100-point increase in the real and nominal interest rates gives a 0.5% reduction of the house price the following year.[31] The relatively high adjusted-R^2 (ranging between 0.32 and 0.38) implies that house price changes seem to be predicatable at a two-year horizon.

Column 4 of Table 3.7 shows the results from pooled regressions with the contemporaneous changes of the same set of factors. As expected, the latter terms are statistically significant and improve the regression fit (the adjusted-R^2 is 0.447). Whereas all the signs and magnitudes of the parameters associated with lagged changes are not much affected, the contemporaneous real disposable income change coefficient is positive and quantitatively very high (0.62). Besides contemporaneous negative effects of nominal interest rate changes, a puzzling positive parameter of real interest rate changes is found. This result could be consistent with households interpreting higher real long term interest rates today as the beginning of further rises in the future, making more profitable to buy a house today rather then to wait. An alternative explanation, which is related to the fact that the change in the nominal interest rate has a negative coefficient of equal size, implying a negative coefficient on the change in inflation expectations, is that agents might expect higher future nominal interest rates depressing their housing demand today.

The results shown in column 5 are robust to country-specific intercepts or fixed effects lowed (column 6). Although the fit worsens slightly (the adjusted-R^2 is 0.438), it is worth pointing out that the Wald test strongly rejects the null hypothesis of a common intercept at any level of significance.

[31] The estimated results are also consistent with a semi-elasticity equal to 1. This can be easily seen by re-

The above specifications include lags of the dependent variable among the regressors. Even if the data set is characterised by a relatively large number of time periods (*T=32)*, there is a vast literature pointing to the fact that fixed effect estimators might be biased and inconsistent in small samples. The robustness of the above results has been tested by applying the Generalised Method of Moments (GMM) Estimator suggested by Arellano and Bond (1991) available in STATA 6. Results shown in the last column of Table 3.7 are remarkably similar to the fixed effects estimates displayed in column 5. On the basis of these findings, it is evident that the inconsistency and bias problems typically found in fixed-effect dynamic panels with large N and fixed T are not a matter of concern for the sample used in this section. Additionally, the sensitivity of the estimates has been checked with the exclusion of countries (namely, Portugal and Austria) with a relatively short time span of the house price index. The main results remain unaffected and are not statistically different from the full sample estimates.

writing eq. $\Delta ph = -0.005\Delta i - 0.005\Delta r$ as $\Delta ph = -0.01\Delta i + 0.005\Delta\pi$ or $\Delta ph = -0.01\Delta r - 0.005\Delta\pi$.

TABLE 3.7 Regression of Real House Price Change on the Change of Explanatory Variables

	1	*2*	*3*	*4*	*5*	*6*
Constant	0.000	-0.004	-0.004	-0.013***	*Fixed Effects*	
DRDI				0.621***	0.657***	0.622***
DRDI(-1)	0.233**	0.301***	0.301***	0.237**	0.312***	0.336**
DRHP(-1)	0.502***	0.599***	0.592***	0.563***	0.541***	0.545***
DRHP(-2)		-0.232***	-0.190***	-0.144***	-0.140***	-0.121**
DRLT				0.004***	0.005***	0.006***
DRLT(-1)	-0.005***	-0.006***	-0.005***	-0.004***	-0.004***	-0.003*
DLT				-0.004**	-0.005**	-0.004
DLT(-1)			-0.005***	-0.006***	-0.006***	-0.007***
DLT(-2)			-0.005**	-0.004**	-0.004***	-0.006**
DDEM(-1)		0.517*	0.464			
Diagnostics						
Adj. R-squared	0.320	0.356	0.382	0.447	0.438	n.a.
S.E. of regression	0.068	0.064	0.063	0.060	0.060	n.a.
DW	1.671	1.866	1.934	1.972	1.972	n.a.
Countries	13	13	13	13	13	13
Observations	336	324	324	327	327	314
Sample Period	70-01	70-01	70-01	70-01	70-01	70-01

Notes: All regressions are estimated with SUR, with the exception of the last column. The dependent variable is the change of the real house price index (DRHP). DRDI is the change of the real household disposable income. DRLT is the change of the real long-term interest rate. DLT is the change of the nominal long-term interest rate. DDEM is the fist difference of the ratio of the population aged 25-44 over the total. Column 5 allows for different country-specific intercepts or fixed effects. Column 6 shows the results of the GMM Estimator suggested by Arellano and Bond (1991). (***), (**) and (*) show statistical significance at 1, 5 and 10 percent levels.

5.2 *A SUR-ECM and PMG Approach of House Price Determination*

The above analysis shows that residential prices are driven by and can be modelled with a relatively small number of determinants. However, when applying pooled techniques, it is not possible to ignore the risks associated with the imposition of the homogeneity of the parameters (Pesaran and Smith, 1995; Pesaran, *et al.*, 1999a; Baltagi and Griffin, 1997). The above literature argues that the use of standard pooled estimators is subject to substantial bias when the parameters are heterogeneous across countries. Given the high degree of asymmetry characterising the European housing and mortgage systems, this point seems particularly relevant.

An additional concern arises from the non-stationarity features of the variables used in the regressions. In macro panels with large N and large T, the order of integration or cointegration of the variables becomes important and extending the estimation and testing procedures for integrated and co-integrated series to panels is a natural development (Shin, 2001).

On the basis of the above considerations, the order of integration of the variables is directly accounted for and the homogeneity assumption of the regression parameters formally tested within an unrestricted Error Correction Mechanism (ECM) approach. The latter is estimated through the standard SURE method and the Pooled Mean Group (PMG) estimator recently developed by Pesaran, *et al.* (1999a). In particular, consistently with the unit root tests carried out and shown in Appendix 2,[32] a

[32] For each annual series the Phillips and Perron (PP) and the Augmented Dickey Fuller (ADF) tests over the period 1970-2001 are implemented and the results shown in Appendix 2. With the exception of very few cases, results indicate that the real house price index (RHP), the real household disposable income (RDI) and the real and nominal interest rates (RLT and LT) are all integrated of order one. The unit root tests for the demographic ratio DEM are very mixed, not only across countries, but are also dependent upon the specific test applied. As a result, this section does not include this variable in the cointegrating relationship.

long-run relationship between the real house price (RHP), the real household disposable income (RDI) and the pre-tax real long term interest rate (RLT) is assumed.[33] Contemporaneously, short-run dynamics are captured by lagged changes of real house prices, contemporaneous and lagged changes of real disposable income, and real and nominal long term interest rates.

In particular, a "general to specific" selection approach is carried out according to which statistically insignificant terms are deleted. The starting general specification included country-specific intercepts[34] and at least two lags of the explanatory variables regressors. Moreover, all the short and long-run coefficients are restricted to be the same across countries. In particular, the selection approach starts from the following unrestricted ECM specification:

$$(14) \qquad \Delta y_{it} = \mu_i + \sum_{j=1}^{2} \lambda_j^{*} \Delta y_{i,t-j} + \sum_{j=0}^{2} \Delta \mathbf{X}_{i,t-j} \boldsymbol{\delta}_j^{*} + \phi\, y_{i,t-1} + \boldsymbol{\beta}' \mathbf{X}_{i,t} + \varepsilon_{it}$$

where $t = 1,2,\ldots,T$ and $i = 1,2,\ldots,N$, μ_i represent the country-specific intercept or fixed effect, $\mathbf{X}_{i,t}$ is a $1 \times k$ vector of regressors, and λ_j^{*}, $\boldsymbol{\delta}_j^{*}$ are scalar and $k \times 1$ vector of parameters. The final specifications are then selected according to the statistical significance of the relevant parameters.

Table 3.8 shows the main results of the SUR estimation. Columns 1 and 2 provide the estimated parameters of the selected ECM specification with and without

[33] Although simple, a house price equation with a long-run relationship between real house prices, real earnings and real long-term interest rates is also used in the Bank of England Macroeconometric Model (MM) (Bank of England, 2003). The only difference is that, due to the fact the model can become explosive, in their house price equation they impose a long-run income elasticity equal to one. The latter assumption, however, is not consistent with a number of other estimated house price equations where the plausible elasticity could range between 0.7 and 2.5 (Meen and Andrew, 1998a and Meen, 2001). This point is further discussed below.

[34] The *F* and *chi-square* Wald test statistics (7.1 and 93.1 respectively) reject the null hypothesis of common constant term in favour of fixed effects.

Germany, Austria and Portugal, respectively. The latter two countries have been excluded because of the relatively short sample of the house price indices (only 12 annual observations), whereas Germany is the only country where BIS data are very different from the ones collected by the ECB. As displayed in the table, however, point estimates are very similar in both specifications. The only exception is the long-run parameter associated to the real long-term interest rate, which becomes statistically significant at five percent level only when excluding the three countries. The adjustment coefficient ranges between -0.16 and -0.17 and the implied long-run income elasticity fluctuates between 0.64 and 0.84. Statistically significant positive short-run effects are given by contemporaneous and lagged changes of real disposable income and real house prices. Negative effects of real and nominal interest rate changes are also found.

Columns 3 and 4 show the results of the same two specifications with the inclusion of a common linear trend in the cointegrating relationship. This is motivated by the need to account for supply effects (e.g. housing stock dynamics) on house price movements. Given that the housing stock tends to grow gradually over time, a linear trend could be a good approximation.[35] This seems to be supported by the fact that this term is found to be strongly significant and with the expected negative sign (columns 3 and 4). Moreover, in both specifications the adjusted R^2 slightly improves. There are, however, changes in the value of the long-run parameter associated with the real disposable income. In particular, a long-run elasticity of 1.27 and 1.70 respectively is

[35] In the countries where some measure of annual housing stock is available (the Netherlands, Denmark, Ireland, the U.K. and Sweden), a constant growth pattern is found. The correlation coefficient with a time trend is in all cases above 0.99. Therefore, a positive linear trend can be assumed to be a good approximation of housing stock dynamics. Kennedy and Andersen (1994) use the same proxy. It is, however, worth pointing out that this term may also capture trends in the quality of dwellings which is not accounted for in same of the house price data.

now found. This is consistent with the fact that the linear trend not only captures supply effects, but also the (upward) trend in disposable income. Moreover, the real long-term rate becomes statistically insignificant in both specifications. All the other short-run parameters change only marginally. The only difference is given by the lagged change of real income and the contemporaneous change in the real long-term rate which are no longer significant. The adjustment coefficients are increased and range between -0.185 and -0.203.

The above findings suggest that the inclusion of a linear trend proxying housing stock dynamics is supported by the data. However, the F and *Chi-square* statistics testing the equality of this term across countries is firmly rejected at any level of significance. As a result, the full sample system is re-estimated with country-specific intercepts and linear trends (column 5). The resulting specification gives a higher adjusted-R^2 and statistically significant long-run parameters associated with real income and real long-term interest rates. The corresponding elasticity and semi-elasticity are now 1.17 and -0.008. Columns 6 and 7 further generalise the latter specification, first by allowing for country-specific long-run income elasticities only, then by allowing both long-run factors to differ across countries. While the common short-term terms are basically unaffected, the equality restrictions on the long-run coefficients are rejected at any level of significance. These results might be dependent upon the assumption of common short-run dynamics. The latter hypothesis is relaxed in column 8, in which the system is estimated only imposing the same error correction term (or adjustment coefficient) and the two long-run parameters, while keeping all deterministic terms and the short-run parameters to be freely estimated. The results show a increase of the adjusted-R^2 (0.53) and statistically significant long-run parameters associated with the real income (1.1), and the real interest rate (-0.013). The

Wald test for the null hypothesis of common short-term dynamics parameters is strongly rejected at any statistical level.

From the above step by step analysis, it is evident that the homogeneity of the regression parameters is always rejected in favour of heterogeneous country-specific effects. To reinforce even further the problems associated with the equality assumptions also in the presence of different specifications for each country, an alternative estimation approach is now applied.

Although in the sample the number of countries N is smaller than the number of observations T, SUR estimation might not be feasible when N tends to T (Pesaran and Shin, 1995). With both N and T sufficiently large, two extreme approaches to estimating dynamic panels have been proposed by the literature.

On one side are the traditional pooled models (e.g. the fixed and random effects, the instrumental variables or the GMM estimators) where only the intercept is allowed to differ across groups, whereas all the other coefficients and error variances are constrained to be the same. Pesaran and Smith (1995) show that, when the slope coefficients differ significantly across groups, these procedures can produce inconsistent and very misleading estimates. As shown in the previous sections, however, a high degree of heterogeneity seems to be present and the use of pooled panels is not supported by the relevant tests. On the other side, the second approach implies taking advantage of the number of observations available for each country by estimating separate equations for each group and pooling only some of the coefficients believed to be similar across countries. Such estimators allow for the imposition of weaker homogeneity assumptions with subsequent efficiency and consistency gains.

TABLE 3.8 SUR - ECM Model for the Real House Price

	1	*2*	*3*	*4*	*5*	*6*	*7*	*8*
				Short-Run Coefficients				
Time trend			-0.003***	-0.004***	*-0.004*	*-0.003*	*-0.006*	*-0.003*
DRDI	0.556***	0.622***	0.600***	0.662***	0.576***	0.457***	0.540***	*0.548*
DRDI(-1)	0.296***	0.190**	0.204**					
DRHP(-1)	0.557***	0.562***	0.553***	0.545***	0.586***	0.585***	0.574***	*0.493*
DRLT	0.003**		0.003**					
DRLT(-1)	-0.003***	-0.002*	-0.004***	-0.004***	-0.003**	-0.004***	-0.003***	*-0.006*
DLT(-1)	-0.006***	-0.007***	-0.006***	-0.007***	-0.005***	-0.004***	-0.004**	*0.001*
RHP(-1)	-0.164***	-0.176***	-0.185***	-0.203***	-0.288***	-0.303***	-0.342***	-0.273***
RDI(-1)	0.106***	0.152***	0.233***	0.342***	0.337***	0.309***	*0.465*	0.294***
RLT(-1)	-0.001	-0.002**	0.001	-0.001	-0.002**	*-0.004*	*-0.006*	-0.002**
				Long-Run Coefficients				
Real Income	0.64	0.84	1.27	1.70	1.17	1.01		1.08
Real Int. Rate	-0.006	-0.012	0.002	-0.004	-0.008			-0.013
Diagnostics								
Adj.R^2	0.482	0.479	0.485	0.494	0.509	0.529	0.529	0.534
S.E.	0.059	0.061	0.059	0.060	0.058	0.057	0.057	0.056
DW	1.987	1.988	1.956	1.970	1.997	2.078	2.017	1.819
Countries	13	10	13	10	13	13	13	13
Observations	336	284	336	284	336	336	336	336
Sample Period	70-01	70-01	70-01	70-01	70-01	70-01	70-01	70-01
Wald Tests H_0:								
Common constant	93.34	88.23	90.58	87.23				
(p-value)	(0.00)	(0.00)	(0.00)	(0.00)				
Common trend			72.61	81.92				
(p-value)			(0.00)	(0.00)				
Common RDI(-1)							34.23	
(p-value)							(0.00)	
Common RLT(-1)						68.26	33.61	
(p-value)						(0.00)	(0.00)	
Common SR Dyn.								103.4
(p-value)								(0.00)

Notes: All regressions are estimated with SUR. The dependent variable is the change of the real house price index (DRHP). DRDI is the change of the real household disposable income. DRLT is the change of the real long-term interest rate. DLT is the change of the nominal long-term interest rate. All equations allow for different country-specific intercepts. Columns 2 and 4 do not include Germany, Austria and Portugal. Column 5 includes country-specific intercepts and linear trends. Columns 6 and 7 generalise the previous specification by allowing for country-specific long-run income elasticities only (column 6) and for both long-run factors to differ across countries (column 7). In column 8 the system is estimated imposing the same error correction term (or adjustment coefficient) and the two long-run parameters, while keeping all deterministic terms and the short-run parameters to be freely estimated across countries. (***), (**) and (*) show statistical significance at 1, 5 and 10 percent level. Values in *italics* are simple group means. The Wald Test tests for the specified null hypothesis are based on the *Chi-square* statistics. *p-values* are in brackets.

Recently, Pesaran, Shin and Smith (1999a) (thereafter PSS) suggested the use of the Pooled Mean Group (PMG) estimator within an ECM framework, which allows the intercepts, time trend, short-run coefficients and error variances to differ freely across groups, while imposing one or more of the long-run parameters to be the same. In particular, they extend the single time series Autoregressive Distributed Lag (*ARDL*) modelling pioneered by Davidson, *et al.* (1978) to the dynamic panel data model as in the following general $ARDL(p,q,...q)$ model:

$$(15) \qquad y_{it} = \alpha_i + \sum_{j=1}^{p} \lambda_{ij} y_{j,t-j} + \sum_{j=0}^{q} \mathbf{x}_{i,t-j} \boldsymbol{\delta}_{ij} + \varepsilon_{it}$$

where $t = 1,2,...,T$ and $i = 1,2,...,N$, α_i is the fixed effect, $\mathbf{x}_{i,t}$ is a $1 \times k$ vector of regressors, and λ_{ij}, $\boldsymbol{\delta}_{ij}$ are scalar and $k \times 1$ vector of parameters. The model in eq. (15) is then re-parameterised in the following unrestricted error correction form:

$$(16) \qquad \Delta y_{it} = \mu_i + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \Delta \mathbf{x}_{i,t-j} \boldsymbol{\delta}_{ij}^{*} + \phi_i y_{i,t-1} + \mathbf{x}_{i,t} \boldsymbol{\beta}_i + \varepsilon_{it}$$

where $\phi_i = -(1 - \sum_{j=1}^{p} \lambda_{ij})$ are the equilibrium (or error correction) parameters and $\boldsymbol{\beta}_i = \sum_{j=0}^{q} \boldsymbol{\delta}_{ij}$ are the long-run parameters.[36] The PMG estimator is based on the assumption that all (or part) of the long-run coefficients are the same across the groups, i.e.

$$(17) \qquad \boldsymbol{\beta}_i' = \boldsymbol{\beta}', \quad i = 1,2,...,N$$

[36] This re-parameterisation differs from traditional unrestricted error correction model pioneered by Sargan (1964) and Davidson, *et al.* (1978) and from the specification estimated in the previous SUR system. In the latter, the long-run coefficients are associated to the lagged levels of the explanatory variables. It can be shown, however, that the two formulations produce identical estimates of the long-run parameters.

and that the $ARDL(p,q,...q)$ model is stable. This implies that equation (16) can be written as:

$$(18) \quad \Delta y_{it} = \mu_i + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \Delta \mathbf{x}_{i,t-j} \boldsymbol{\delta}_{ij}^{*} + \phi_i y_{i,t-1} + \mathbf{x}_{i,t} \boldsymbol{\beta} + \varepsilon_{it}$$

PSS estimate the above equation by maximum likelihood (ML) and test for the long-run equality restrictions applying the likelihood ratio (LR) test.[37] Given the short sample for the Austrian and the Portuguese house price index, the latter two are not included in the dynamic heterogeneous panel, restricting the total number of countries to eleven over the period 1970-2001.

The approach is based on starting from a common general $ARDL(2,2,2)$ for each country separately with a constant, a linear trend, and the current and lagged change of the nominal long-term interest rate as exogenous variables. The specific lag order for each country-specific equation is then chosen according to the Schwarz Bayesian Criterion (SBC). As expected, the selected *ARDL* models differ from country to country, indicating that imposing the same short-run dynamic terms as in the SUR specification shown in Table 3.8 might not be correct. In most cases two lags of the real house price index are chosen, whereas between one or two lags of the real income and real long term interest rates are generally selected. Alternative selection criteria (e.g. the Akaike Information Criterion, and the Hann and Quin Criterion) produce very similar results. The PMG approach is based on the maximisation of the likelihood function, where alternative common long-run parameters could be used as

[37] This estimation is implemented in a GAUSS procedure, which is available at Pesaran's website (http://www.econ.cam.ac.uk/faculty/pesaran/jasa.exe). This approach has been used by a number of authors amongst which Haque, *et al.* (2000), Cameron and Muellbauer (2001) and Ludwig and Sloek (2002) are recent examples.

initial values. The results obtained below, however, are very robust and not affected by the particular starting values used.

Table 3.9 shows the main findings, with the estimated average of the heterogeneous short-run dynamic terms and the common long-run coefficients. Column 1 provides the full sample estimates. Given the relatively low number of countries and the different lag length applied, the short-run coefficient estimates are not representative for all countries and the corresponding sign and statistical significance (based on simple standard deviations of each parameter across countries) are only indicative. The long-run coefficients are very similar to the ones obtained in column 8 of Table 8. In particular, the house price elasticity with respect to real disposable income is 1.11, whereas the semi-elasticity of the real interest rate is -0.011. Both are statistically significant at one percent level. In column 2, Germany is excluded from the sample. Results are very similar, although there is a reduction in the income elasticity to 0.933. The average size of the group-specific error correction coefficients is statistically significant in both samples with values ranging between -0.23 and -0.26.

Although the above results are remarkably similar to the ones obtained in the SUR estimation with country-specific short-run parameters (column 8, Table 3.8), the likelihood ratio test for the homogeneity assumption of the long-run coefficients is strongly rejected at any level of significance. Column 3 (column 4) shows the estimation results imposing the same parameter associated to the level of the real income (real interest rate), leaving the one related to the real interest rate (real income) to be country specific. The relevant LR tests both reject the homogeneity assumption.

Columns 5 and 6 show the results of the PMG estimation for a sub-set of countries. Namely, the sample is split into two groups, according to *a-priori* beliefs on the relative similarity of the factors characterising the specific housing and mortgage systems across Europe (e.g. the competitiveness of banking sector, the degree of financial liberalisation, the level of leverage of households, etc.). Although subjective, according to the above criteria, Group 1 includes the United Kingdom, Sweden, Finland, Denmark and the Netherlands, whereas Group 2 is formed by the remaining countries, namely Italy, France, Germany, Ireland, Spain and Belgium.[38] The results are very interesting. While the income elasticities are similar across the two groups (0.925-0.912, the long-run semi-elasticity associated with the real interest rates is very high and statistically significant only in Group 1 (-0.028), and insignificant and close to zero in Group 2. Column 7 excludes Germany from Group 2, but results are not affected in a significant way. What is worth pointing out, however, is that in all regression, the LR test rejects the equality of the single long-run parameters at one percent level.

Baltagi and Griffin (1997) observe that even though formal tests for equality of the parameters across country are rejected, most researchers proceed to estimate pooled models. Pesaran and Smith (1995) and Pesaran, Shin and Smith (1999a), however, show that the presence of heterogeneous parameters across countries could lead to seriously biased coefficients. From the SUR and the PMG estimation analysis above, it has been shown that the homogeneity of the regression coefficients appears to be a very poor assumption. This is not surprising, given the high degree of heterogeneity in the housing and mortgage systems across Europe. In what follows, these asymmetries are directly accounted for by estimating individual country regressions where not only

[38] This classification is also consistent with the results of chapter 2.

the short-run dynamic specification is allowed to differ, but also the long-run parameters and the adjustment coefficients are country specific. At the same time, however, in order to enhance comparability and interpretation of the estimation results, the same methodology and similar specifications are used.

TABLE 3.9 Pooled Mean Group (PMG) Estimation for Real House Price

	1	*2*	*3*	*4*	*5*	*6*	*7*
			Averages of Heterogeneous Short-Run Coefficients				
Constant	-0.106***	0.077***	-0.139***	0.082	0.077***	0.151***	0.026
Time trend	-0.001	0.001	-0.002	-0.002	-0.002	-0.002	0.000
DRDI	0.318*	0.228	0.182	0.465	0.228	0.268	-0.001
DRDI(-1)	0.106	0.009	0.080	-0.214	0.009	0.157	0.157*
DRHP(-1)	0.522***	0.563***	0.568***	0.766***	0.563***	0.563***	0.408**
DRLT	0.003**	0.004**	0.006**	0.005**	0.009**	0.004*	0.003
DRLT(-1)	-0.001	-0.002	-0.001	-0.000	-0.000	-0.001	-0.001
DLT	-0.001	-0.008	0.000	-0.002	-0.008	-0.001	0.006
DLT(-1)	-0.005*	-0.002*	-0.004*	-0.005**		-0.005**	-0.005***
Adjustment coefficient	-0.260***	-0.233***	-0.306***	-0.299***	-0.304***	-0.233***	-0.30***
			Long-Run Coefficients				
Real Income	1.113***	0.933***	1.132***	*0.941***	0.925***	0.912***	0.984***
Real Int. Rate	-0.011***	-0.013***	*-0.013**	-0.011***	-0.028**	-0.000	-0.004
Diagnostics							
LR Test	99.50	60.04	65.53	61.87	25.90	44.44	32.01
(p-value)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Countries	11	10	11	11	5	6	5

Notes: All regressions are estimated with the Pooled Mean Group (PMG) Estimator method suggested by Pesaran *et al.* (1999a). The dependent variable is the change of the real house price index (DRHP). DRDI is the change of the real household disposable income. DRLT is the change of the real long-term interest rate. DLT is the change of the nominal long-term interest rate. Column 1 reports the results of the full sample. Column 2 excludes Germany. Column 3 (column 4) shows the estimation results imposing the same parameter associated to the level of the real income (real interest rate), leaving the one related to the real interest rate (real income) to be country specific. Column 5 shows the results for Group 1 (the United Kingdom, Sweden, the Netherlands, Denmark and Finland). Column 6 refers to Group 2 (Italy, France, Spain, Ireland, Belgium and Germany). Results in *italics* are mean group estimates. The last column excludes Germany from Group 2. The Likelihood Ratio (LR) test tests for the equality restriction of the long-run coefficients. (***), (**) and (*) show statistical significance at 1, 5 and 10 percent level.

6. Individual House Price Equations

6.1 Data

This section uses quarterly data for eleven of the 15 EU countries.[39] With the only exception of Italy, Portugal and Spain, where the sample period starts in 1984, 1987 and 1988 respectively, in most of the remaining countries the equations are estimated from the early 1980s to the beginning of 2001. The advantage of using quarterly data in contrast to annual data is that there are more observations with subsequent efficiency gains and a better estimation of short-run dynamics. The disadvantage, however, is that for some variables (i.e. the demographic variables) there are no quarterly series, and interpolation techniques have to be implemented. Moreover, for some countries more than ten years of valuable information contained in annual data is lost with the high frequency series.

The variables chosen in these specifications are partly based on previous house price specifications estimated by the literature and partly constrained by the information available at European level. The ultimate goal of this section is not to fit the data in the best possible way for forecasting exercises, as several individual country studies do, but to find a simple specification, which can help to disentangle the role of some real and financial variables during the 1980s and 1990s, possibly controlling for country-specific events which occurred in the mortgage and housing markets.

For each country under investigation, the following variables are been collected and constructed: the real house price index (RHP), the real household disposable income (RDI), the nominal mortgage rate (MRATE), the real mortgage rate

[39] Results for Germany, Luxembourg, Greece and Austria are not provided because the series are either too short, or the number of observations too small. For Italy the house price index is available at semi-annual frequency only. As a result, the estimation is implemented with this frequency.

(RMRATE), two mortgage market condition variables (that is, the real mortgage stock RMSTOCK and the ratio of the nominal mortgage stock over the nominal GDP, MSTOCKY) and the ratio of the population aged 25-44 over the total (DEM). Although different cohorts might be more suitable in North European countries, a broader population proportion could be a better proxy to pick up demographic effects on housing demand in Continental housing systems, where the average age for buying a house is relatively higher. Hort (1998), for instance, finds that this ratio captures this effect on the Swedish housing demand relatively well. A measure of the real user cost variable (UC), defined as the real mortgage rate minus expected housing capital gains (proxied by the real house price inflation over the last year) is also constructed.[40] A detailed description of the sources and definitions of the data is provided in Appendix 1.

6.2 *Econometric Methodology*

The vast majority of the housing literature on the U.K. and Continental Europe applies dynamic equation approaches or unrestricted single equation error correction model (ECM), which were first advocated by Sargan (1964) and Davidson, *et al.* (1978), and more recently appreciated in Benerjee, *et al.* (1993, 1998), Pesaran and Shin (1999) and Ericsson and MacKinnon (2002). The latter authors discuss the advantages and disadvantages of this approach,[41] arguing that the ECM procedure is robust to many particulars of the marginal process, that is, the specific lag lengths and dynamics

[40] The backward-looking expectations assumption is discussed below. As discussed above, a more accurate measure of the real user cost should include the country specific income and property tax rates, as well as maintenance costs and depreciation rate as in eq. (10). Such information is very difficult to collect or not available for most of the countries under investigation. Real mortgage rates and housing capital gains, however, are quantitatively the two most important (and variable) components of the real user cost.

[41] This approach can be validly used only if there is a single cointegrating vector and the regressors are assumed to be weakly exogenous. To verify whether the former assumption is validated by the data, the Johansen approach for a subset of countries has been implemented. Results show strong evidence in support of the existence of a unique cointegrating relationship in the two specifications. Details of the main results can be found in Appendix 3.

involved.[42] They also provide critical values for the single equation error correction statistic for testing cointegration.

A general formulation of an unrestricted ECM model can be given by the following equation:[43]

$$(19) \quad \Delta y_t = \beta_0 + \sum_{i=1}^{p-1} \beta_{1,i} \Delta y_{t-i} + \sum_{j=0}^{q-1} \boldsymbol{\beta}_{x,j} \Delta \mathbf{x}_{t-j} + \alpha (y_{t-1} - \boldsymbol{\varphi} \mathbf{x}_{t-1}) + \varepsilon_t$$

where $\mathbf{x}_t$ is a vector of n weakly exogenous variables assumed to co-integrate with y_t, α is the error-correction coefficient determining the speed of adjustment to the long-run equilibrium, $\boldsymbol{\varphi}$ the vector of long-run coefficients, and ε_t a white noise error term. Direct estimation of eq. (19) not only allows us to obtain the long-run parameters and the corresponding adjustment term, but also the short-run dynamics, which are given by contemporaneous and lagged changes of the variables $\mathbf{x}_t$ and lagged changes of the endogenous variable y_t. Additional exogenous variables could be included on the right hand side to capture specific short-run dynamics.

To select the final house price specification, a 'general-to-specific' modelling approach is implemented. In particular, the initial broad over-parameterised model contains at least five lags of the short-run dynamics variables. The best specification is then selected on the basis of the Akaike Information Criterion (AIC) and the Schwarz Bayesian Criterion (SC), which weigh the goodness of fit of a model together with the

[42] An additional reason for using a single equation approach over the multivariate Johansen cointegration methodology is due to the use of interpolation to obtain quarterly data and, as a result of that, the fact that it is extremely difficult to find an adequate reduced form for the demography equation. Pain and Westaway (1996, 1997) find similar difficulties. Moreover, an unrestricted estimation method is chosen over a traditional two-step Engle-Granger technique because the latter might be affected by small sample bias (Banerjee *et al.*, 1993). Robustness tests of the single-equation ECM results with respect to a typical two-stage Engle Granger approach show very similar results.

[43] The model below can be considered as a re-parameterisation of a general autoregressive distributed lag $ARDL(p, q, ..q)$ model. Chapter 8 of Johnston and DiNardo (1997) provides a comprehensive outline of this approach.

number of estimated parameters. During this selection process, a series of diagnostic tests for autocorrelation, heteroskedasticity, normality and parameter stability are implemented. In particular, if the specification survives these tests, the strategy is to delete the statistically insignificant variables only if both information criteria improve.

6.3 Order of Integration of the Variables

An important step before implementing any cointegration analysis is to assess their order of integration. Appendix 2 shows the results of the unit root tests in the variables used in the house price models: namely, the real house price index (RHP), the real household disposable income (RDI), the real and nominal mortgage rates (RMRATE and MRATE)), and the proxy for the real user cost of housing (UC). In particular, they all appear to be non-stationary in levels, but stationary in first differences, regardless of the deterministic terms and the method adopted. The only exception is France, where the real user cost appear to be I(0). As for the two mortgage condition proxies (MSTOCKY and RMSTOCK), they are found to be first-differences stationary.[44] Finally, the demographic ratio (DEM) is integrated of order one in Belgium, Finland, Denmark and the United Kingdom. In all the others countries, except Sweden and Spain (where it is stationary in level), DEM appear to be I(2). Consistent with the unit root tests, the latter variable might enter the cointegrating vector in levels or first differences. Due to the low power of the above unit root tests, alternative orders of integration are assumed and tested for in the cointegration analysis below.

[44] It is only in the U.K. that the presence of a unit root in the second difference of MSTOCKY cannot be rejected. The latter variable is therefore assumed to be integrated of order two.

6.4 *The 'User-Cost' House Price Equation*

Following the recent literature based on the life-cycle model outlined above and simplified in equation (13), this section show the results of a house price equation in which some proxy of the user cost of housing is directly included. This calls for the construction of a measure of the expected housing capital gains. In two influential papers, Muellbauer and Murphy (1997) and Hendry (1984) argue that households do not possess all the relevant knowledge of the mortgage and housing markets. Therefore, they assume 'semi-rational' and some form of 'sensible' expectations, where households are expected to have just some general information on interest rates, mortgage availability, past values of house prices and so on.

Other authors (Di Pasquale and Wheaton, 1994; Clayton, 1998, Meen, 1998, 2000, 2002), on the other hand, argue that housing cycles are fairly predictable and reject housing market efficiency due to slow adjustments to demand and supply changes. Following the latter point, most of the empirical studies assume an extrapolative behaviour, and model expectations about future house prices as backward-looking processes. Meen (2001), Pain and Westaway (1996) discuss the role of expectations in the housing market literature, showing that this is an assumption generally supported by the data.[45] A direct investigation of house price expectations is also provided in Case and Shiller (1988) and Nordvick (1995). The latter study points out that "*price expectations are formed through an extrapolation of the trend in house prices rather than through a process based on knowledge of the structure of the housing market, economic fundamentals and demographic trends.*" Similarly, Case and Shiller

[45] An alternative approximation would be to use the actual house price inflation rate over the allowing year and, due to the presence of error-in-variables problems, estimate the models with instrumental variables. While the results are qualitatively similar to the ones shown in this section, the parameters seem to be very sensitive to the instruments chosen.

(1988) argue that rising prices in the near past generate expectations of rising prices in the near future. As a result, in this chapter a backward-looking or adaptive expectations are adopted and the expected housing capital gains are proxied with the actual house price inflation rate over the last year.[46]

The user-cost house price specification estimated in this section assumes the existence of a long-run relationship between the real house price (RHP), the real household disposable income (RDI), the real user cost of housing (UC)[47] and the demographic ratio (DEM).[48] In section 3 it has been argued that the household financial wealth might be included in the cointegrating vector to proxy permanent income, which is considered to be determinant of the unobservable imputed rental price of housing services. Financial wealth data, however, are not readily available for all the countries under investigation. Therefore, to maintain a similar general specification, the marginal effect of household financial wealth is not accounted for. All variables, but the real user cost, are in logs. Similar to the panel estimation of the previous section, a time trend is included in the general specification, implying the imposition of a linear trend in the cointegrating vectors. As discussed above, this approximation ought to capture housing stock effects.[49]

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[46] This assumption is less controversial as long as no bubble develops. If, however, the house prices deviate too far from fundamentals more and more people might become rational, making the extrapolative hypothesis less attractive. I would like to thank Jan-Egbert Sturm for raising this point.

[47] Given the limited availability of mortgage interest rates for Italy and Portugal, the lending rate to the private sector and the long-run interest rate, respectively, are used. Over the overlapping period the two series are highly correlated with the corresponding mortgage rates. Similarly, the Netherlands do not have a reasonable (and representative) measure of the typical mortgage rate (see Appendix 1). As a result, the lending rate to the private sector is used.

[48] Following the results of the unit root tests in Appendix 2, DEM enters in level in Belgium, Finland, the U.K. and Denmark, and in first difference in all the other countries, except Spain and Sweden. As a robustness check, alternative orders of integrations of the demographic ratio are also assumed including the latter variable in the cointegrating vector in levels and first difference.

[49] In a recent review of the theoretical and empirical literature, Meen and Andrew (1998a) argue that the most important variables quantitatively in explaining house prices are income, interest rates, demographic structure and the housing stock. The chosen specification contains all these factors.

The French real user cost is found to be stationary in levels, and from a theoretical point of view could not cointegrate with the other variables. Given the low power of the unit root tests and assuming the possibility that UC is integrated of order one, the existence of a cointegration vector which includes the latter variable has been tested. Results, however, show no evidence in support of a stable cointegrating vector. In Portugal, probably due the short sample period and few degrees of freedom, the resulting house price equation parameters appear to be very unstable with economically implausible signs. Consequently, the results for either of the two countries are not reported.

The short-run dynamics are modelled by including lags of the first difference of the dependent variable (DRHP), and contemporaneous and lagged values of the change of the real disposable income (DRDI), and the real and nominal mortgage rate (DRMRATE and DMRATE). The significance of the latter variable is tested to account for the presence of 'front-end-loading' effects discussed above. Finally, to control for financial liberalisation effects and relaxation of credit constraints, lagged changes of the ratio of nominal mortgage stock to nominal GDP (DMSTOCKY) or the real mortgage stock (RMSTOCK) are included.

Before implementing the unrestricted ECM approach, the existence of a unique cointegrating vector was tested by using the Johansen (1988, 1995) methodology. Appendix 3 shows the main results with reference to a subset of countries (namely Finland, Belgium and Sweden). Results show that the trace and eigenvalue test statistics cannot reject the existence of a single cointegration vector. The resulting long-run parameters, however, are not always qualitatively plausible, and are very much dependent upon the imposed deterministic terms, the number of lags chosen and the dummies that are necessary to obtain normal residuals in all the equations of the unrestricted VAR systems. Because of these problems and degree of freedom considerations, similarly to the vast majority of the housing literature, a single unrestricted ECM method seems feasible.

The general specification starts with an unrestricted fifth order ECM model, where no allowance is made for short-run demographic effects. As in Pain and Westaway (1997), given the need for interpolation to obtain quarterly data and the risk of spurious results, these short-run dynamics terms are not included. Indeed, if included, they have either no predictive power, or, when statistically significant, have economically implausible coefficients. In the latter case, however, by excluding them, both the significance and the sign of the other terms of the dynamic equation, and the fit of the equation were not affected. Therefore, also on the basis of the assumption that demographic effects are expected to have mainly long-run influences, they are not included as short-run dynamic terms.

To improve the precision of the estimates, obtain well-behaved residuals and take into account country-specific events which occurred in the national housing and mortgage markets (namely major reforms in taxes and credit markets), a series of dummy variables has been included.[50] In all specifications, however, these dummies do not much affect the size and the significance of the short- and long-run parameters. A constant and a set of seasonal dummies have also been included.

Table 3.10 shows the final data-based restricted ECM specifications which are selected following the general-to-specific approach discussed above. In all equations, a

[50] In most cases they refer to particular tax breaks (i.e. 1983 and 1991 in Sweden, 1988 in the United Kingdom) and significant mortgage market developments (1992 in the Netherlands and Spain, 1991 in Portugal) which, when available, have also been documented by the relevant housing literature.

battery of diagnostic tests to check for the presence of serial correlation, heteroskedasticity and non-normality in the residuals have been carried out during and after the selection of the best specification. On the basis of the relevant statistics, all equations seem to pass the tests at any significance level. The only exception is found in Spain, where the White's test statistic could not reject the null hypothesis of no heteroskedasticity. The standard errors are then corrected by using the heteroskedasticity consistent covariance matrix (White, 1980). Similarly, CUSUM tests suggest parameter stability in all equations. The Ramsey's test for the functional form is rejected at five percent level only in Sweden. The adjusted R^2 is very high, ranging between 0.59 in Ireland and 0.93 in Italy.

The signs of the estimated long-run coefficients are the expected ones. In particular, with only the exception of Spain, which shows a very high income elasticity (around 4), in all the other countries it ranges between 0.72 in Ireland and 2.03 in Italy and Denmark. The size of the latter figures is consistent with a number of previous empirical studies based on the U.K, the U.S. and Scandinavian countries. Meen and Andrew (1998a) argue that, whereas from a purely theoretical point of view it might be implausible to have a long-run income elasticity of house prices in excess of unity, the fact that in the national house price literature they systematically range between 1.5 and 3, is consistent with a low price elasticity of housing demand.[51]

[51] This point could be appreciated by looking at a typical house price equation estimated as an inverted housing demand function given in eq. (1). Assuming a simplified version, e.g. $H_t^d = \alpha_0 + \alpha_1 ph_t + \alpha_2 y_t$, the elasticity of house prices with respect to income is given by the ratio of the income elasticity of housing demand (α_2) to the price elasticity (α_1). The empirical evidence summarised in Meen and Andrew (1998a) shows that in the U.K., the U.S. and Sweden the former ranges between 0.5 and 1, whereas the latter between -0.4 and -0.7. As a result, feasible estimates of the income elasticity of house prices are in the range 0.7-2.5. On the basis of the above considerations, very low price elasticity of housing demand could lead to income elasticity of house prices well above 2.5-3. See Mean and Andrew (1998a) for a full discussion.

Additionally, it is not clear how different degrees of financial liberalisation across countries could be reflected in the sensitivity of house prices to income fluctuations. While there is an extensive literature based on the *excess sensitivity* of consumption to income shocks which might imply lower income elasticities of house prices as credit constraints are relaxed, recent studies (Lamont and Stein, 1999 and Almeida, 2001) have shown that housing demand and house prices are an increasing function of housing finance development.[52] Therefore, without a full theoretical model it is difficult to provide a reasonable economic interpretation of the income elasticities found above.

As for the long-run coefficients associated with the real user cost of capital, with the exception of Italy, statistically significant negative semi-elasticities, which differ from country to country, are found. For instance, in Denmark, Sweden, the Netherlands and the United Kingdom, a permanent increase in the level of the real user cost by 100-points (say from 4% to 5%) produces a permanent reduction in the *level* of real house price by 2.5-4%. In the remaining countries, the same shock determines a permanent drop of real house prices by about 0.7-1.9%. The elasticity associated to the demographic variable is found to be statistically significant only in Finland.[53] Finally, the time trend significantly enters the cointegrating vector in two countries, Belgium and Spain, but it has the expected negative sign only in the latter.

As far as the short-run dynamics are concerned, in all countries but Finland, changes in the nominal mortgage interest rates have a strong predictive power and the corresponding parameters are all negative. This is consistent with the recent house price booms, which according to many authors might well have been fuelled by the gradual

[52] See the above references for a full explanation of this relationship.
[53] The significance of demographic factors in Finland is found also by Kostela *at al.* (1992).

reduction of nominal lending rates and a greater housing lending affordability. In Denmark, Finland, the U.K., and Sweden either lagged changes of the ratio of nominal mortgage stock over nominal GDP or the real mortgage stock are statistically significant and have the expected positive effects. The fact that these variables have a predictive power only in those countries where the financial liberalisation process has been more pronounced indicates that, although not ideal, this proxy might capture periods of credit easing.

Finally, the coefficient associated with the level of the real house price (that is, the error correction or adjustment coefficient) has the expected negative sign and is statistically significant at 1% level in all countries. In particular, it ranges between a relatively low -0.06 in the United Kingdom and -0.20 in Belgium. This indicates that about 6-20 percent of the disequilibrium from the estimated long-run relationship is absorbed within a quarter.

Ericsson and MacKinnon (2002) provide critical values for testing whether the level variables in an error correction equation are co-integrated. This implies comparing the estimated *t*-statistic correspondent to the level of the real house price with their critical value, which depends on the number of variables and the specific deterministic terms. The five percent critical values for the case of an intercept and no trend in the unrestricted ECM model and four variables in the cointegrating vector is -3.75. In the case of a constant and a linear trend the critical value is -4.14. With three variables the corresponding values are -3.50 and -3.92. The above unrestricted ECM specification contains seasonal dummies and additional exogenous variables.[54] As a result, the above

[54] Pesaran *et al.* (1999) report alternative critical values under the assumption that the underlying variables are either I(1) or I(0). In particular, they compare the *F*-test for the joint significance of the lagged level variables in the unrestricted ECM model with an 'uncertainty band'. Our results are consistent with the presence of a cointegrating vector for most of the countries.

tests are not directly applicable. Nevertheless, in all countries except Spain, the relevant t-statistics are well above the appropriate critical values providing some additional support in favour of a long-run cointegrating relationship.[55]

[55] The t-statistics in Belgium, Spain and Sweden (where a trend and a constant are included) are 6.22, 3.50 and 4.71, respectively. In the remaining countries (where only the constant is included) the t-statistics are all above 4.02.

TABLE 3.10 The 'User-Cost' House Price Equation. Dependent Variable: Change of Real House Prices (DRHP)

	BE	*DK*	*ES*	*FI*	*GB*	*IE*	*NL*	*IT*	*SE*
Short-Run Coefficients									
Constant	-0.280*	-0.418**	-2.257***	0.313***	-0.169**	0.202*	-0.293***	-0.678**	-0.023
Time Trend	0.001***		-0.006***						0.000
DRHP(-1)	-0.609***			0.396***		-0.369**	-0.545***	0.429**	-0.313**
DRHP(-2)	-0.353***	-0.210*				-0.597***	-0.412**	0.859***	-0.056
DRHP(-3)	-0.295***	-0.311***				-0.132	-0.328*		-0.183
DRHP(-4)		-0.249**				0.174	-0.272		
DRHP(-5)		0.196***				0.378***			
DRDI	0.805***	1.057***	0.612**	0.470**	0.046	0.590*	0.770***	0.300	0.316*
DRDI(-1)		-0.106	-0.310	-0.796***	-0.147				-0.483**
DRDI(-2)		-0.691***	-0.592**		0.095				-0.084
DRDI(-3)		0.706***			0.002				-0.179
DRDI(-4)					0.265**				-0.453**
DMRATE		-0.025***			-0.008***				
DMRATE(-1)	-0.007*		0.013*		-0.005**	-0.023***			
DMRATE(-2)	-0.007*		-0.013*			0.021***	-0.011***	-0.007***	-0.009***
DMRATE(-3)						-0.018***			
DMSTOCKY(-1)		0.530***							
DRMSTOCK(-1)	0.117			1.077***	0.639***				0.304***
DRMSTOCK(-2)					0.528***				
Long-Run Coefficients									
RHP(-1)	-0.201***	-0.097***	-0.190***	-0.160***	-0.056***	-0.142***	-0.127***	-0.137***	-0.094***
RDI(-1)	0.228***	0.197***	0.873***	0.174***	0.101***	0.102***	0.195***	0.278***	0.153**
UC(-1)	-0.002***	-0.003***	-0.002***	-0.001***	-0.001***	-0.003***	-0.005***	0.002	-0.004***
DEM(-1)				0.295***					

TABLE 3.10 The 'User-Cost' House Price Equation. Dependent Variable: Change of Real House Prices (DRHP) (continued)

	BE	*DK*	*ES*	*FI*	*GB*	*IE*	*NL*	*IT*	*SE*
Implied LR parameters									
Real Household Disposable Income	1.134***	2.030***	4.593***	1.087***	1.800***	0.720***	1.535***	2.036***	1.762**
Real User Cost	-0.011***	-0.030***	-0.010***	-0.007***	-0.025***	-0.019***	-0.037***	0.016	-0.040***
Demographic				1.848***					
Diagnostics									
Adjusted R-squared	0.824	0.780	0.783	0.792	0.810	0.587	0.749	0.934	0.845
S.E. of regression	0.011	0.015	0.011	0.016	0.012	0.019	0.016	0.013	0.010
Durbin-Watson stat	1.857	2.140	1.839	1.941	2.126	2.201	1.633	2.153	1.914
Breusch-Godfrey LM Test	0.85	0.43	1.24	0.39	0.28	1.04	1.5	1.51	0.88
Normality Jaque-Bera's Test	0.91	0.65	0.29	2.27	0.44	0.78	5.29*	3.05	0.63
White Heteroskedasticity Test	0.92	1.25	3.05***	1.47	1.12	0.63	0.99	1.23	0.91
Ramsey's RESET	0.84	0.94	1.68	0.41	0.18	1.68	0.81	0.25	2.50*
CUSUM	x	x	x	x	x	n.a.	x	x	x
Dummies	82Q3 82Q4	93Q2	92Q1 92Q2	90Q2 91Q2	88Q3 92Q2	91Q2	92Q2	89S1 87S2	83Q1 92Q1
Observations	72	79	47	81	79	66	77	36	78
Adjusted Sample Period	82Q2-00Q1	80Q3-00Q1	88Q2-00Q1	81Q1-01Q1	81Q3-01Q1	82Q4-99Q1	81Q1-00Q1	84S1-01S2	80Q4-00Q1

Notes: For variables definitions see the text and Appendix 1. All regressions include seasonal dummy variables which are deleted if not statistically significant. (*), (**) and (***) denote statistical significance at 10, 5 and 1 percent level. The Breusch-Godfrey Lagrange Multiplier test shows the *F*-statistic for the null hypothesis of no up-to-the fourth-order serial correlation in the residuals. The normality Jarque-Bera's test is distributed as a *Chi-square* distribution with two degrees of freedom under the null of normal residuals. The White's Heteroskedasticity test shows the F-statistic and the null is the presence of no-heteroskedastic residuals. Under the presence of heteroskedasticity, the White heteroskedasticity consistent covariance estimates are applied. Ramsey's RESET is the F-statistic for testing the hypothesis that the coefficients on the powers of fitted values are all zero. An "x" in the CUSUM of squares test box indicates that the plot of the relevant CUSUM statistic moves within the 5 percent critical lines, suggesting parameter stability. An "n.a." stands for not applicable.

6.5 *The 'Reduced-Form' Demand-Supply House Price Equation*

In the housing empirical literature, some measure of the cost of housing financing is generally included in the estimated house price equations. As in many British and Scandinavian studies, in the previous section mortgage interest rates effects are accounted for through a proxy for the real user cost of housing. Given the way that the latter variable is constructed, however, it is hard to provide a direct estimate of house price elasticity with respect to the real cost of housing lending. Moreover, due to the extrapolative expectation assumption made to measure housing capital gains, it is not clear whether the statistically significant negative effects associated with the user cost are due to real interest rate fluctuations or house price capital gains. With persistent house price changes, the extrapolative assumption might hide the 'pure' effect of real mortgage interest rate changes on housing demand. Additionally, from a policy-maker's perspective, it is more informative knowing the long-run elasticity of house price with respect to a direct measure of real lending rates.[56]

This section estimates an alternative and, maybe, more 'transparent' house price equation, which has been widely used in the academic literature (see the seminal papers by Hendry, 1984; Nellis and Longbotton, 1981; Meen, 1990 and, more recently, Kasparova and White, 2001). In particular, similarly to the specification suggested by the demand-supply approach in equation (3), this section tests for the existence of a long-run relationship between the real house price (RHP), the real household disposable income (RDI), the real mortgage interest rate (RMRATE)[57] and the same demographic

[56] As emphasised in section 5, the Bank of England Macroeconometric Model (MM) includes a house price equation where the real long-term interest rate is included in the cointegrating vector. To our knowledge, the UK model is the only central bank macro-model explicitly including a house price equation.

[57] The real mortgage rate is defined as the difference between the nominal mortgage rate and the actual HICP annual inflation rate.

variable used in the previous specification (DEM).[58] As in the user cost model, a linear trend is included to proxy housing stock dynamics. The mortgage stock dynamics are not included in the cointegrating vector in that they are endogenous to real house price fluctuations. Given the weakly exogeneity assumption required by the single equation approach, similarly to the user cost specification, the effects of mortgage lending availability are captured by lagged changes in the ratio of nominal mortgage stock to nominal GDP (DMSTOCKY) or real mortgage stock (RMSTOCK) which are pre-determined variables. Additional short-run dynamics are given by lags of the first difference of the dependent variable, and contemporaneous and lagged changes of the RDI, and the real and the nominal mortgage rates. In all equations, the diagnostic tests for serial correlation, heteroskedasticity and normality of the residuals, and functional form and parameter stability do not indicate any particular problem.[59] The only exception is found in Spain, where the Ramsey's test is significant at 1% level.

As in the previous specification, all the estimated long-run coefficients associated to the real disposable income have the expected signs and their sizes are very similar to the ones shown in Table 3.10. In France and Portugal, which were not included in the user-cost specification, the long-run income elasticities are 4 and 1.3, respectively. The coefficients related to the real mortgage interest rate are statistically significant and negative in all countries but Italy and Belgium. The corresponding semi-elasticity is in the range of a relatively low -0.015 in Portugal and more then -0.033 in Finland, Denmark, the Netherlands, the United Kingdom and Sweden. This implies that in the latter countries a reduction of real mortgage rate by 100-points (say from 5% to 6%) generates a 3.3-4% increase in the level of real house prices. Besides Finland, the

[58] As in the previous section, Appendix 3 shows that the assumption of a single cointegrating vector is supported by the data, making the single equation estimation an adequate approach.

demographic variable is only statistically significant in France. Finally, the time trend appears to be statistically significant in five countries. With the exception of Belgium, the corresponding parameter is negative.

As for the short-run dynamics, in all countries the nominal mortgage rate changes still have a strong predictive power and, with the exception of Portugal, the coefficient shows the expected negative sign. At least one of the two measures of financial liberalisation is found positive and statistically significant in the same four countries (Denmark, Finland, the U.K. and Sweden) as the user-cost specification.[60] The adjustment coefficients are remarkably similar to the user cost model. The only exception is found in the Netherlands, where the error correction term is now much smaller (from 0.12 to 0.07).[61]

Overall, from the results of the two models, it seems that the two specifications might well be regarded as useful tools to study the determinants of house prices across European countries. While more complex specifications, accounting for additional country-specific factors, might improve the fit each individual house price model, the relatively parsimonious estimated models are, however, very informative. What seems interesting from the above cross-country analysis in relation to the previous international empirical evidence is the fact that housing and mortgage markets are very diverse across the EU and that the imposition of the same structure for empirical purposes might produce misleading results. In fact, both the typical user cost approach

[59] In a few cases, namely Spain, France and Ireland, AR terms were included to obtain white noise residuals.
[60] Some effect is also found in Portugal.
[61] Comparing their t-statistics with the 5% critical values suggested by Ericsson and MacKinnon (2002), it is found that, with the exception of Portugal, Spain and the Netherlands, in all the remaining countries, the hypothesis of no cointegration is rejected.

and the more *ad-hoc* reduced form specification indicate that over the last twenty years the same driving factors have had different effects on residential prices in Europe.

Although the specific coefficient values might be affected by the presence of heterogeneous definitions of the national house price indexes and the mortgage market variables (i.e. interest rates and stocks), results are consistent with a greater sensitivity of house prices to interest changes in the U.K., Denmark, Sweden, Finland and the Netherlands. For countries like Belgium and Italy, real variables (namely, the real disposable income) seem to be the main determinant. In the case of the remaining countries (Spain, France, Ireland and Portugal) it is hard to provide a clear-cut classification. Nevertheless, it is encouraging that results are very much in line with the findings of chapter 2, which shows an asymmetric sensitivity of house prices to temporary short-term interest rate shocks.

Moreover, it is found that the proxies for financial liberalisation developments in the mortgage market have a strong predictive power in those countries that witnessed a more intense deregulation process. Additionally, in almost all specifications and countries, nominal mortgage rate changes seem to have played a significant role, which is consistent with the recent house price rises experienced in the vast majority of the European countries during the last few years.

TABLE 3.11 The 'Reduced-Form' House Price Equation. Dependent Variable: Change of Real House Prices (DRHP)

	BE	DK	ES	FI	FR	GB	IE	NL	IT	SE	PT
Short-Run Coefficients											
Constant	-0.160	-0.287	-1.741***	0.249***	-0.661	-0.284***	0.236	-0.225**	-0.686**	-0.059	
Time Trend	0.001***		-0.006***		-0.004***					-0.001*	-0.001***
DRHP(-1)	-0.401***	0.285***	0.544***	0.457***	0.144			-0.066	0.184***	0.107	
DRHP(-2)	-0.149		-0.054	0.249***	0.303***	0.139*	-0.255**	0.064	0.610***	0.373***	-0.267***
DRHP(-3)	-0.072		0.154	0.092	0.423***	0.010	0.156	0.211**		0.238***	-0.234***
DRHP(-4)	0.207**		0.175	0.151*	0.261***	0.255***	0.414***	0.292***		0.426***	0.177**
DRHP(-5)		0.203***	-0.360***				0.305***				
DRDI	0.798***	1.139***	0.946***	0.474**	-1.413**		0.636*	0.628**	0.257	0.209	
DRDI(-1)		-0.427*	0.079	-0.754***						-0.353*	
DRDI(-2)		-0.286	-0.446*							-0.208	0.473***
DRDI(-3)		0.612***	0.588**							-0.254	0.326***
DRDI(-4)						0.426***				-0.493**	
DRMRATE								0.010**			
DMRATE	0.007*	-0.026***		-0.010*		-0.008***		-0.009*			
DMRATE(-1)							-0.022***	-0.005			0.006***
DMRATE(-2)	-0.010**		-0.010**				0.023***	-0.010***	-0.007***	-0.009***	
DMRATE(-3)					-0.010***		-0.018***				
DMSTOCKY(-1)										0.282***	
DRMSTOCK(-1)		0.678**		1.108***		1.185***					
DRMSTOCK(-2)						0.549***					
DRMSTOCK(-3)		0.350*									0.341***
Long-Run Coefficients											
RHP(-1)	-0.204***	-0.130***	-0.158***	-0.192***	-0.210***	-0.059***	-0.160***	-0.066**	-0.139***	-0.108***	-0.094***
RDISI(-1)	0.196***	0.204***	0.734***	0.322***	0.859***	0.131***	0.114***	0.117**	0.283***	0.189**	0.121***
RMRATE(-1)	-0.001	-0.005**	-0.005**	-0.006***	-0.004*	-0.002**	-0.004*	-0.003*	0.002	-0.004***	-0.001**
DEM(-1)				0.661***	1.310***						

TABLE 3.11 The 'Reduced-Form' House Price Equation. Dependent Variable: Change of Real House Prices (DRHP) (continued)

	BE	*DK*	*ES*	*FI*	*FR*	*GB*	*IE*	*NL*	*IT*	*SE*	*PT*
Implied LR parameters											
Real Income	0.960***	1.560***	4.631***	1.678***	4.079***	2.190***	0.711***	1.763***	2.020***	1.744**	1.280***
Mortgage Rate	-0.002	-0.040**	-0.029**	-0.033***	-0.020*	-0.038**	-0.023*	-0.040*	0.015	-0.038***	-0.015**
Demographic				3.440***	6.223***						
Diagnostics											
Adjusted R-squared	0.816	0.786	0.842	0.812	0.873	0.820	0.568	0.750	0.934	0.841	0.751
S.E. of regression	0.011	0.015	0.009	0.015	0.011	0.012	0.020	0.016	0.013	0.011	0.006
Durbin-Watson stat	1.832	2.091	2.277	1.997	1.918	1.837	2.135	1.842	2.111	1.912	1.841
Breusch-Godfrey LM Test	0.81	0.39	1.02	0.25	1.20	1.39	0.63	0.62	1.02	0.88	0.22
Normality Jaque-Bera's Test	0.85	0.62	2.01	2.19	0.23	0.2	0.56	2.98	3.13	0.46	1.81
White's Heteroskedasticity Test	1.28	0.65	1.39	1.16	1.00	0.81	0.54	1.41	1.05	1.31	1.40
Ramsey's RESET	1.09	0.29	8.39***	0.93	0.48	0.49	1.73	0.96	0.13	2.17*	0.43
CUSUM	x	x	n.a.	x	n.a.	x	n.a.	x	x	x	x
Dummies	82Q3 82Q4	93Q2	92Q1 92Q2	90Q2 91Q2	82Q2 85Q3-4	88Q3 92Q2	91Q2	92Q2	89S1 87S2	83Q1 92Q1	91Q4
Observations	72	79	47	81	79	79	66	77	36	78	44
Adjusted Sample Period	82Q2-00Q1	80Q3-00Q1	88Q2-00Q1	81Q1-01Q1	82Q2-01Q4	81Q3-01Q1	82Q4-99Q1	81Q1-00Q1	84S1-01S2	80Q4-00Q1	89Q2-01Q1

Notes: For variables definitions see the text and Appendix 1. All regressions include seasonal dummy variables which are deleted if not statistically significant. (*), (**) and (***) denote statistical significance at 10, 5 and 1 percent level. The Breusch-Godfrey Lagrange Multiplier test shows the F-statistic for the null hypothesis of no (up-to-the fourth-order) serial correlation in the residuals. The normality Jarque-Bera's test is distributed as a *Chi-square* distribution with two degrees of freedom under the null of normal residuals. The White's Heteroskedasticity test shows the F-statistic and the null is the presence of no-heteroskedastic residuals. Under the presence of heteroskedasticity, the White heteroskedasticity consistent covariance estimates are applied. Ramsey's RESET is the F-statistic for testing the hypothesis that the coefficients on the powers of fitted values are all zero. An "x" in the CUSUM of squares test box indicates that the plot of the relevant CUSUM statistic moves within the 5 percent critical lines, suggesting parameter stability. An "n.a." stands for not applicable.

7. *Conclusions*

This chapter studies the determinants of national house prices over the last few decades in most of the European countries. The first part provides a detailed descriptive analysis focusing on the degree of volatility, and the nature of the booms and busts in the housing sector across countries over the period 1980-2001. Some of the main developments which occurred in the mortgage markets during the financial deregulation process of the 1980s and 1990s are also discussed. Besides specific individual features related to the housing tax treatment, transaction costs, variable-versus-fixed mortgage interest rates and the typical maturity of mortgage contracts, additional factors related to the developments of nominal and real interest rates over the same period are also analysed. From the above discussion, it emerges that, besides common macroeconomic factors (i.e. interest rates dynamics during the pre-EMU period and financial deregulation process), European housing and mortgage systems are still characterised by a high degree of heterogeneity which might imply a different sensitivity of house prices to real and financial shocks.

The above conclusion is formally tested by extending the previous international empirical literature modelling house prices using the annual BIS data set over the period 1970-2001. In particular, pooled and heterogeneous panel estimation techniques are implemented and the equality assumption of the coefficients associated with the factors driving residential prices are formally tested for. After systematically rejecting the latter hypothesis, the following sections make use of a higher frequency data base, which contains quarterly information on national house price indexes, real household disposable income, outstanding mortgage stock, typical mortgage interest rates and demographic variables over the period 1980-2001. To account for the most important short and long-run determinants of real house prices, an unrestricted error correction mechanism (ECM) approach is implemented. From the estimated models, it is found that real disposable income is the most import factor affecting house prices, although the corresponding long-run elasticity differs from country to country. Additionally, European national house prices seem to be asymmetrically affected by financial variables, partly reflecting the institutional features of their housing and mortgage systems. In particular, results indicate that real mortgage rates constitute a relatively more influential long-run factor for housing demand in those countries where the housing finance markets are relatively more liberalised and competitive (namely the U.K., Denmark, Finland, Sweden and the Netherlands). While demographic factors are statistically significant in only few countries (France and Finland), nominal mortgage interest rate changes are found to have a very strong predictive power in all estimated specifications across Europe, indicating that their recent fall might have contributed to the house price booms of the 1990s.

Overall, although the results might be partly affected by a low comparability and quality of housing statistics, the main findings are supportive of the presence of a high degree of heterogeneity in European housing markets. While stronger market integration and growing financial deregulation within the euro-area will probably increase the efficiency of the housing and mortgage markets and its degree of competition, EU countries are still very diverse in their respective legal systems as well as culturally. These aspects seem to be the cause of a diverse sensitivity of house prices to macroeconomic shocks. In particular, from a monetary policy perspective, the fact that the national housing markets are asymmetrically affected by interest rate dynamics implies that the housing channel enters the MTM heterogeneously across European countries. These conclusions are consistent with the results of chapter two.

APPENDIX 1 *Data Description and Sources*

This Appendix describes the definition and the sources of the data, partly provided by the National Central Banks (NCBs), and partly collected from the ECB, IFS, OECD, AMECO, BIS, EMF and European Commission databases.

GROSS DOMESTIC PRODUCT at constant prices. The quarterly real GDP series is from ECB and is based on ESA 95. With the exception of Austria, Italy, Portugal and Ireland, the original series are seasonally adjusted with the X-12 multiplicative method available in E-views 4.

REAL HOUSEHOLD DISPOSABLE INCOME. The annual real household disposable income series are from the OECD Economic Outlook. The quarterly frequency was provided by Torsen Sloek of the IMF.

CONSUMER PRICE INDEX. The series for the EMU-12 countries are the harmonized consumer price index (HICP) produced by the ECB. Before 1995 the series are estimated by the ECB. The series for Sweden, Denmark and the United Kingdom are from national sources. The correspondent codes are BISM.m.VEBA.SE.01, BISM.m.VEBA.DK.02 and BISM.m.VEBA.GB.01.

LONG-TERM INTEREST RATES. The nominal annual long-term interest rates (LT) are from the European Commission database (AMECO). The code is AME.A.....1.3.0.0.ILN. The quarterly series are from the DERIVED database of the ECB with the code Q....IRL.AV.Z. For Finland, Ireland, Sweden and Denmark data are provided by the NCB. In the U.K. the Main Economic Indicators (MEI) series is used (MEI.Q.GBR.IRLTGV02.ST). The real long-run interest rates have been calculated as the difference of the nominal rates minus the HICP inflation rate over the last year.

HOUSE PRICE INDEX. This chapter uses different house price data sets. For annual frequencies the official BIS database, whose sources and descriptions are shown in Table A.1.1, is taken. As for the higher frequency series collected by the ECB, the relevant information is outlined in Table A.1.2. Although the two datasets are not homogeneous, the only country where the dynamics of the two indexes are different is Germany.

In both tables, data refer to transaction prices recorded by private real estate associations and, in many cases, are collected and published by central banks and central statistical offices in different countries. House price developments are at national level and are generally based on sales prices of existing dwellings or, when not available, those of new dwellings. The only exceptions are France and Austria, where data refer to the area of their respective capitals, Paris and Vienna. Differences in the construction of these indices, the characteristics of dwellings (i.e. mix of properties traded over time, quality changes, etc) and the composition of the dwelling stock within and across countries render quantitative attempts to draw strong international comparisons less robust.

Real house price are constructed as nominal house prices deflated by the harmonised consumer price index. For Denmark, Sweden and the United Kingdom the standard consumer price index (CPI) is used.

TABLE A.1.1 BIS Annual House Prices in the EU

	Availability	*Source*	*Description*
Belgium	1970-2001	Antwerpse Hypotheek Bank	Price index for small and middle houses
Denmark	1970-2001	Danmarks Statistik	Residential property prices: index of cash price of one-family houses "purchase price at cash value in % of officially appraised cash value in 1992"
Finland	1970-2001	Property Advisors Helsinki	Index of price/m2 of old housing corporation flats in deals brokered by real estate agents
France	1970-2001	Ministère de l'Equipement, du Logement et des Transports	Sales prices of new apartments in France
Germany	1970-2001	Bundesbank	Price of terrassed houses (Western Germany)
Ireland	1970-2001	Investment Property Databank	Average prices of new houses for which loans were approved by "All Agencies"
Italy	1972-2001	Nomisma, Osservatorio sul Mercato Immobiliare; Bologna	House price index, weighted average of new, second hand and refurbished dwellings
Netherlands	1970-2001	Statistics Netherlands, CBS	Sales prices of existing flats
Spain	1975-2001	Ministerio de Obras Publicas, Transportes y medio ambiente (Banco de Espana, Boletin Estadistico)	Residential property prices per square meter
Sweden	1970-2001	Statistics Sweden; Monthly Digest	House price index; owner occupied one and two-dwelling buildings
United Kingdom	1970-2001	Department of the Environment, Transport and the Regions (DETR)	UK mix-adjusted house price index, all dwellings; based on survey of all lenders

Sources: Bank for International Settlements (BIS) (using national data).

TABLE A.1.2 ECB Higher Frequency House Prices in the EU

	Availability from	*Source*	*Definition/Notes*
Belgium	81Q1	Antwerspe Hypotheek Bank	Price index for small and middle houses
Denmark	71Q1	National Statistical Office	Cash price of one-family houses sold. Nationwide.
Finland	70Q1	Bank of Finland	Average price for existing flats per square meter
France	80Q2	INSEE, based on Chambre des Notaires data	Average price per square meter for existing dwellings sold in Paris
Ireland	78Q1	Department of the Environment and Local Government	Average gross prices of new (and second-hand) houses for which loans were approved by "All Agencies"
Italy	72S1	Il Consulente Immobiliare	Residential property index
Netherlands	77Q1	NVM	Average Selling Price of existing single and multy-family houses (50-60% of total transactions)
Portugal	88M1	Confidencial Imobiliare	New and existing dwellings
Spain	87Q1	INE – Ministerio de Economia y Hacienda	Residential Property Price Index per Square Meter
Sweden	72Q1	Central Statistical Office (CSO)	House Price Index as weighted mean of primary and leisure homes
United Kingdom	74Q2	Department of the Environment, Transport and the Regions (DETR)	UK mix-adjusted house price index, all dwellings; based on survey of all lenders

TABLE A.1.2 ECB Higher Frequency House Prices in the EU (continued)

	Availability from	*Source*	*Definition/Notes*
Austria	87S1	Wiener Immobilienbörse/ SRF-TU Wien	New and existing dwellings in Vienna
Germany	80A	National Central Bank	Aggregate compiled from series for new and existing flats and for new and existing terraced houses.
Greece	93Q4	National Central Bank	Residential property prices (chained index, initially covering 13 cities, later extended to 15 cities outside Athens, NSA)
Luxembourg	74A	National Central Bank	New and existing one family houses

Sources: European Central Bank (ECB). With only few exceptions (i.e. France and Germany), the definition of higher frequency house price indexes coincides with the one of the BIS data set. The latter, however, does not include Portugal, Austria and Luxembourg.

LENDING RATES TO THE PRIVATE SECTOR. The lending interest rate refers to the average lending rate applied to the private sector and is available in the IFS. The code is IFS.Q….60P..ZF...

OUTSTANDING STOCK OF HOUSING LENDING. The series on the nominal value of the stock of housing lending to households are provided by the NCBs. Definitions of each series and methods used to construct are not homogeneous across countries, but they are within the same national housing mortgage markets. Data are available starting from 1980.

POPULATION. Annual data have been collected from New CRONOS of the European Commission and linearly interpolated to obtain quarterly frequency. The above data have been used to construct the ratio of the population between 25 and 44 over the total to account for demographic factors.

MORTGAGE INTEREST RATES. These have been collected from two main sources: NCBs and European Mortgage Federation (EMF). Although an effort has been made to obtain a representative mortgage interest rate weighted according to the typical mortgage instruments offered in each country, in some cases, due to availability, they refer to a specific contract. As a result, it is not advisable to make cross-country comparisons concerning the level of these interest rates. In some cases, in order to extend the time availability, annual data are interpolated using the growth rate of the above long-term interest rates. Table A.1.3 shows the sources and, when available, the definitions. The real mortgage interest rates (RMRATE) are constructed using the consumer price index inflation over the last year. The real user cost of housing services (UC) is defined as real mortgage interest rate, minus the real house price inflation rate over the last year.

TABLE A.1.3 Mortgage Interest Rates in the EU

	Definition and Construction Methods
DE	Average of three MFI mortgage interest rates for residential sites: floating interest rate, 5 year fixed rate and 10 year fixed rate.
FR	Weighted average of fixed (80%) and variable (20%) interest rates on mortgages.
BE	Up to 1992Q4 average of mortgage rate (redeemable through constant depreciation; 20 years rate; renewable after 5 years) and mortgage rate (redeemable through mixed life insurance; 20 years rate; fixed rate). Thereafter, this rate has been linked to the rate communicated to the ECB, which is constructed with the 20 years rate and the renewable after 5 years rates (on the basis of a survey driven by NCB amongst 25 credit institutions, insurance corporations and mortgage credit companies, weighted by the outstanding credit).
ES	Average interest rate on new mortgage loans to households for house purchase (in domestic currency) offered by commercial and saving banks.
NL	Nominal 5-years rate on mortgages for residential real estate and shop-property.
GB	Average effective mortgage rate.
IE	Variable Mortgage Rates - All Mortgage Lenders. Midpoint, I.e. Average of lowest and highest rates.
PT	Link in 1991 of the mortgage rate published by the EMF (weighted average of all lenders) and the Portuguese contribution to the overall monetary union rate on housing loans to households.
SE	Average of the daily Sweden 5-year (Spintab) mortgage lending rate and the Stadshypotekets lån till villor löptid 5 år (mortgages to villas, 5 year).
FI	Average of mortgage interest rates to the households referred to new loans and the stock.
AT	Up to 1995 the EMF re-viewable rate. Thereafter Austrian contribution to the overall monetary union rate on housing loans to households.
DK	Definition not available.
GR	Definition not available.
IT	Definition not available.

Sources: European Central Bank (ECB).

APPENDIX 2 *Unit Root Analysis*

This Appendix provides the results of the unit root tests. In particular, for each of the variables used in the chapter, both the Phillips-Perron (PP) and the Augmented Dickey-Fuller (ADF) tests are implemented. Given that in most of the cases the two approaches provide very similar results, the tables below show the PP test results only. When the two methods suggest different orders of integration, both tests are displayed.

ANNUAL DATA

Section 5 makes use of annual series over the sample period 1970-2001. In particular, the main variables are the following: the real house price (RHP), the real disposable income (RDI), the nominal and the real long-term interest rates (LT and RLT) and the population aged 25-44 over the total ratio (DEM). From Tables A.2.1, A.2.2 and A.2.3 it is possible to verify that, with the only exception of few cases (Austria and Portugal), RHP, RDI, LT and RLT all appear to be integrated of order one. The unit root tests for the demographic ratio DEM are very mixed not only across countries, but are also dependent upon the specific test applied. In particular, the PP test indicates that in Ireland, Finland, Denmark and the U.K. DEM is integrated of order one, while all the remaining countries (apart from Portugal) appear to be I(2). With the exception of the Netherlands, Belgium, Spain, Italy and Austria in all the other countries these results are not confirmed by the ADF tests.

TABLE A.2.1 Unit Root Tests for RHP and RDI - Annual 1970-2001

	RHP		*RDI*	
	Level	First Difference	Level	First Difference
BE	-0.82	**-2.37****	-2.75	**-4.45*****
ES	-0.77	**-2.31****	-2.38	**-3.36****
FI	-1.41	**-2.37****	-2.14	**-4.32*****
FR	-1.54	**-2.55****	-3.37	**-3.49****
NL	-0.76	**-2.36****	-2.18	**-4.35*****
GB	-1.26	**-2.40****	-2.28	**-4.70*****
SE	-1.88	**-2.57****	-2.55	**-3.47****
DK	-1.86	**-3.45*****	-2.71	**-4.91*****
IE	0.82	**-2.27****	0.73	**-3.52****
DE	-2.67	**-2.96*****	-2.19	**-3.65****
IT	-2.55	**-4.28*****	-2.55	**-3.76*****
AT	-1.12	-1.27	-2.85	**-3.77*****
PT	-0.94	**-2.84*****	**-4.97*****	-

Notes: PP tests in level include a constant and a trend. For the first-difference tests a constant only is included for RDI, whereas no deterministic term for RHP. The truncation lag is two. (**) and (***) denote significance at 5% and 1% level. The critical values with a trend and a constant are -4.28 and -3.56, with a constant only -3.66 and -2.96, and with no deterministic trend -2.66 and -1.95.

TABLE A.2.2 Unit Root Tests for LT and RLT - Annual 1970-2001

	LT		*RLT*	
	Level	First Difference	Level	First Difference
BE	-1.60	**-3.86*****	-1.69	**-4.37*****
ES	-1.15	**-4.05*****	-1.70	**-5.17*****
FI	-1.15	**-4.37*****	-1.72	**-5.47*****
FR	-1.71	**-4.38*****	-1.54	**-5.57*****
NL	-1.96	**-4.91*****	-1.22	**-5.15*****
GB	-2.09	**-4.44*****	-2.67	**-14.03*****
SE	-0.76	**-5.66*****	-2.50	**-8.26*****
DK	-2.26	**-5.98*****	-1.61	**-6.59*****
IE	-2.69	**-4.93*****	-2.00	**-4.85*****
DE	-2.69	**-4.37*****	-2.25	**-7.25*****
IT	-1.35	**-3.35*****	-1.61	**-4.75*****
AT	-1.95	**-4.59*****	-2.06	**-5.59*****
PT	-0.97	**-2.98****	**-3.55****	-

Notes: PP tests for the nominal (LT) and the real (RLT) long-term interest rate in level include a constant and a trend. In the first difference tests a constant only is included. The truncation lag is two. (**) and (***) denote significance at five and one percent level. The respective critical values with a trend and a constant are -4.28 and -3.56, and with a constant only are -3.66 and -2.96.

TABLE A.2.3 PP and ADF Unit Root Tests for DEM - Annual 1970-2001

	PP			*ADF*		
	Level	First Difference	Second Difference	Level	First Difference	Second Difference
BE	-0.83	-1.95	**-6.86*** **	-2.88	-0.83	**-8.07*** **
ES	-1.78	-1.84	**-5.61*** **	-2.14	-1.78	**-5.60*** **
FI	1.08	**-3.62*** **	-	-0.34	1.08	**-6.08*** **
FR	0.53	-2.16	**-4.55*** **	**-3.16** **	-	-
NL	2.96	-3.33	**-6.05*** **	0.90	2.96	**-6.47*** **
GB	-1.55	**-1.96** **	-	-1.67	-1.55	**-3.43*** **
SE	-1.76	-2.52	**-4.03*** **	**-4.75*** **	-	-
DK	-1.13	**-4.04*** **	-	-2.47	-1.13	**-4.46*** **
IE	-0.87	**-3.18** **	-	-2.43	-0.87	**-8.64*** **
DE	-1.89	-1.70	**-5.34*** **	**-4.28*** **	-	-
IT	-2.65	-1.77	**-5.33*** **	-2.28	-2.65	**-5.11*** **
AT	-2.56	-1.66	**-5.13*** **	-2.66	-2.16	**-5,07*** **
PT	**-4.15** **	-	-	-2.71	**-4.18** **	-

Notes: In the PP test a truncation lag of two was used. The ADF test has been carried out with up to 2 lags to make the residuals white noise. (**) and (***) denote significance at five and one percent level. For each country and variable different determinist terms have been selected.

QUARTERLY DATA

Section 6 makes use of quarterly series over the sample period 1980-2001. The variables (with the respective abbreviations in brackets) are the following: the real house price (RHP), the real household disposable income (RDI), the nominal mortgage rate (MRATE), the real mortgage rate (RMRATE), the real user cost of housing (UC), the population aged 25-44 over the total ratio (DEM), the real housing mortgage stock (RMSTOCK) and the ratio of the nominal housing mortgage stock over the nominal GDP (MSTOCKY).

Tables A.2.4 and A.2.5 show the results of the PP unit root test for the RHP, RDI, RMRATE and MRATE. In all countries, the null hypothesis of the presence of a unit root in the variables in levels cannot be rejected. Taking the first difference, on the other hand, produces quite clear-cut results, rejecting it at 5% significance level. After double-checking with the ADF tests and alternative deterministic terms, it is reasonable to conclude that the above four variables are all I(1).

Results in Table A.2.6 show that the real user cost is I(1), with the only exception of France, where it appears to be stationary in levels. The two mortgage liberalisation proxies MSTOCKY and RMSTOCK are not stationary in levels, but in first differences. In the United Kingdom and Italy, the first variable and the second variable respectively, appear to be I(2).

Finally, given the mixed results obtained with the PP test for the demographic variable, the ADF tests with different deterministic terms are also displayed. By construction, DEM is a ratio which is bounded between 0 and 1. Over the long-run, it might be expected to be stationary in level. However, in the sample period, the PP tests cannot reject the presence of a unit root only in Spain and Sweden. The ADF test, on the other hand, seems to suggest that the level of DEM is stationary only in Italy. The same tests, applied to the first difference, do not reject the null hypothesis of no stationarity in Belgium, Finland, the United Kingdom and Denmark. In all the remaining countries, stationarity is obtained only after differencing twice. From the above results, it is very difficult to 'classify' DEM in any clear-cut manner not only across, but also within countries. Previous empirical papers (Kennedy and Andersen (1994), Kostela *et al.* (1992), Hort (1998), Breedon and Joyce (1993), among others), find (or directly assume) first difference stationarity. Given the low power of the tests, the results of the unrestricted ECM have been tested assuming alternative orders of integrations to the ones found in this appendix.

TABLE A.2.4 **Unit Root Tests for RHP and RDI – Quarterly 1980-2001**

	RHP		*RDI*	
	Level	First Difference	Level	First Difference
BE	-3.42	**-8.94*****	-2.58	**-4.41*****
ES	-2.18	**-3.71****	-2.18	**-3.78*****
FI	-1.60	**-3.72****	-1.40	**-3.74*****
FR	-1.10	**-4.60*****	-1.88	**-3.58*****
NL	-3.45	**-8.19*****	-3.08	**-4.58*****
GB	-1.25	**-7.30*****	-1.64	**-3.11****
SE	-1.82	**-5.43*****	-1.69	**-4.65*****
DK	-1.83	**-5.33*****	-2.45	**-4.59*****
PT	-1.72	**-5.13*****	-1.75	**-10.79*****
IE	-1.50	**-8.52*****	-1.53	**-13.63*****
IT	-1.51	**-4.35*****	-2.10	**-4.01*****

Notes: PP tests for the variable in level include a constant and a trend. For the first differenced variables test a constant only is included. A truncation lag of three was used in the PP test. (**) and (***) denote significance at 5% and 1% level. The respective critical values with a trend and a constant are -4.06 and -3.46, whereas with a constant only are -3.51 and -2.89.

TABLE A.2.5 **Unit Root Tests for MRATE and RMRATE - Quarterly 1980-2001**

	MRATE		*RMRATE*	
	Level	First Difference	Level	First Difference
BE	-2.27	**-6.56*****	-2.49	**-4.41*****
ES	-1.78	**-5.12*****	-1.79	**-3.78*****
FI	-1.48	**-5.13*****	-0.99	**-7.09*****
FR	-3.23	**-7.68*****	-2.67	**-3.58*****
NL	-2.18	**-6.98*****	-2.14	**-4.58*****
GB	-2.41	**-7.89*****	-2.88	**-3.11****
SE	-3.02	**-6.74*****	-3.14	**-4.65*****
DK	-2.25	**-5.85*****	-1.98	**-4.59*****
PT	-1.77	**-3.73*****	-2.33	**-10.79*****
IE	-3.19	**-6.77*****	-1.61	**-13.63*****
IT	-2.31	**-4.29*****	-2.92	**-4.94*****

Notes: PP tests for the variable in level include a constant and a trend. For the first differenced variables test a constant only is included. A truncation lag of three was used in the PP test. (**) and (***) denote significance at 5% and 1% level. The respective critical values with a trend and a constant are -4.06 and -3.46, whereas with a constant only are -3.51 and -2.89.

TABLE A.2.6 **Unit Root Tests for MSTOCKY, RMSTOCK and UC – Quarterly 1980-2001**

	MSTOCKY		*RMSTOCK*		*UC*	
	Level	First Difference	Level	First Difference	Level	First Difference
BE	-1.80	**-8.32*** **	-2.96	**-7.63*** **	-2.73	**-9.31*** **
ES	-3.18	**-6.03*** **	-1.07	**-4.89*** **	-1.66	**-4.94*** **
FI	-1.46	**-3.41*** **	-1.45	**-3.75*** **	-2.28	**-4.11*** **
FR	-1.83	**-6.48*** **	-1.07	**-5.93*** **	**-6.95*** **	-
NL	-0.56	**-3.61*** **	-2.02	**-4.16*** **	-2.83	**-8.44*** **
GB	-1.26	-2.78	-0.80	**-5.00*** **	-2.84	**-6.96*** **
SE	-0.10	**-4.71*** **	-0.22	**-6.16*** **	-2.20	**-6.20*** **
DK	-2.01	**-4.38*** **	-1.47	**-4.27*** **	-2.91	**-6.53*** **
PT	-0.85	**-3.43** **	-0.79	**-4.24*** **	-3.03	**-9.72*** **
IE	-2.06	**-5.58*** **	-2.61	**-5.16*** **	-2.73	**-9.31*** **
IT	-2.32	**-3.03** **	-1.85	-2.45	-1.78	**-4.54*** **

Notes: PP tests for the variable in level include a constant and a trend. For the first differenced variables test a constant only is included. A truncation lag of three was used in the PP test. (**) and (***) denote significance at 5% and 1% percent level. The respective critical values with a trend and a constant are -4.06 and -3.46, whereas with a constant only are -3.51 and -2.89.

TABLE A.2.7 **PP and ADF Unit Root Tests for DEM**

	PP Tests			*ADF Tests*		
	Level	First Difference	Second Difference	Level	First Difference	Second Difference
BE	1.94	**-2.02** **	-	-2.06	**-2.32** **	-
ES	**-4.80*** **	-	-	-1.71	-0.29	**-4.04*** **
FI	-1.18	**-1.98** **	-	-2.12	**-1.97** **	-
FR	1.30	-0.64	**-9.26*** **	-0.27	-0.87	**-3.12*** **
NL	1.67	-1.06	**-9.28*** **	0.56	-0.96	**-4.17*** **
GB	-0.93	**-1.94** **	-	-2.75	**-2.20** **	-
SE	**-4.19*** **	-	-	-3.19	**-2.57** **	-
DK	-3.35	**-2.15** **	-	-2.22	**-2.15** **	-
PT	-0.79	-1.08	**-9.29*** **	-2.94	-0.80	**-5.34*** **
IE	-2.40	-1.29	**-9.30*** **	-2.32	-0.89	**-6.67*** **
IT	-2.30	**-5.60*** **	-	**-3.59** **	-	-

Notes: In the PP test a truncation lag of three was used. The ADF test has been carried out with up to 8 lags to make the residuals white noise. The tests for the variables in level include a constant (C) and a trend (T). For the second difference, tests are carried out with no deterministic terms. In the first difference tests, country-specific determinist terms were selected. (**) and (***) denote significance at 5 and 1 percent level.

APPENDIX 3 The Johansen Cointegration Analysis

This Appendix shows the results of the Johansen (1988, 1995) cointegration analysis for some of the EU countries under investigation, namely Belgium, Finland and Sweden. A multivariate cointegrating approach is implemented because in a system with more than two I(1) variables, there might be multiple cointegrating vectors (CV), which would invalidate the findings of a conventional single equation approach (Harris, 1995).

On the basis of the two models outlined in section 6, the existence of a long-run relationship between the real house prices (RHP), the real disposable income (RDI), the real user cost (UC) (or the real mortgage rate, RMRATE), the demographic factor (DEM) and the stock of dwellings is assumed. Given the lack of available data for all countries under investigation, the latter variable is proxied with a linear time trend in the cointegrating vector. Appendix 2 shows that the population share DEM seems to have heterogeneous univariate features both with annual and quarterly data. As for the higher frequency time series, for Belgium, Finland, the United Kingdom and Denmark, it is found to be integrated of order one, whereas in the other countries, except Spain and Sweden, it seems to be I(2). From the unit root analysis it also emerges that all the other variables, with the exception of France, are integrated of order one. As a result, in the specific sub-set of countries, the level of the demographic factor in the cointegrating vector is only included in Belgium and Finland. Besides quarter specific dummy variables, necessary to ensure that the residuals of the unrestricted VAR models are normally distributed, additional exogenous stationary variables are included to control for financial liberalisation and tilting effects, that is, lags of the first difference of the real mortgage stock (DRMSTOCK) and nominal mortgage rates (MRATE).

Tables A.3.1, A.3.2 and A.3.3 show the Johansen tests of cointegration for Belgium, Finland and Sweden, respectively. For Belgium a four variables system, (including RHP, RDI, RMRATE and DEM) is first estimated. The maximal eigenvalue and the trace tests indicate the existence of one cointegrating vector, but the long-run parameter associated with DEM not statistically different from zero (results are not shown). Therefore, the latter variable is excluded. Table A.3.1 summarises the Johansen tests for the three variable system only. In all specifications, the tests cannot reject the existence of a single cointegrating vector. With the exception of Model 3, in which no time trend is included, the normalised (with respect to the real house price) long-run coefficients have the correct sign with plausible magnitudes. For Sweden, very similar results are found, although the size of the long-run coefficients is very different (Table A.3.3).

Table A.3.2 shows the Finnish results of the Johansen tests in the four variable system. All tests, regardless of the deterministic terms and the alternative cointegrating vectors (with the real mortgage rate or the user cost proxy), cannot reject the existence of a single cointegrating vector. It is, however, worth pointing out that, in order to obtain normally distributed residuals in the reduced form equation of the population share in the unrestricted VAR, several dummies correspondent to the first quarter of each year were included. The same problems were encountered for Belgium. This is not particularly surprising, in that the demographic ratio was obtained by linear interpolation from annual data. Pain and Westaway (1997) report similar difficulties in their analysis applied to the U.K. over the period 1969-93.

Overall, the above results provide strong evidence in favour of the existence of a single cointegrating relationship both in the specification of the user cost and the one of the real mortgage rate in all three countries. Nevertheless, in a few cases the estimated cointegrating vectors produce coefficients which are not economically interpretable, whereas in other cases they are very sensitive to the chosen specification and the lag length. On the basis of the above findings, the difficulty in obtaining adequate reduced-form VAR equations for the variables other than the real house price and a more direct comparability with the existing empirical literature, an unrestricted error correction mechanism (ECM) model is the approach implemented in the chapter.

TABLE A.3.1 Johansen Tests for Cointegration for Belgium

Model 1

Cointegration with unrestricted intercepts and no trends in the VAR.
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	23.99	21.12	19.02
r<= 1	r = 2	**7.51****	14.88	12.98
r<= 2	r = 3	3.58	8.07	6.50

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	35.09	31.54	28.78
r<= 1	r>= 2	**11.10****	17.86	15.75
r<= 2	r = 3	3.58	8.07	6.50

Model 2

Cointegration with unrestricted intercepts and restricted trends in the VAR.
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	34.76	25.42	23.10
r<= 1	r = 2	**17.87***	19.22	17.18
r<= 2	r = 3	5.85	12.39	10.55

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	58.48	42.34	39.34
r<= 1	r>= 2	**23.72***	25.77	23.08
r<= 2	r = 3	5.85	12.39	10.55

TABLE A.3.1 Johansen Tests for Cointegration for Belgium (continued)

Model 3

Cointegration with unrestricted intercepts and no trends in the VAR.
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	**15.37****	21.12	19.02
r<= 1	r = 2	9.10	14.88	12.98
r<= 2	r = 3	0.39	8.07	6.50

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	**24.86****	31.54	28.78
r<= 1	r>= 2	9.49	17.86	15.75
r<= 2	r = 3	0.39	8.07	6.50

Model 4

Cointegration with unrestricted intercepts and restricted trends in the VAR.
Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	34.76	25.42	23.10
r<= 1	r = 2	**17.87***	19.22	17.18
r<= 2	r = 3	5.85	12.39	10.55

Cointegration LR Test Based on Trace of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	**24.86****	31.54	28.78
r<= 1	r>= 2	9.49	17.86	15.75
r<= 2	r = 3	0.39	8.07	6.50

Estimated Normalised Cointegrating Vectors

	Model 1	Model 2	Model 3	Model 4
RHP	1.00	1.00	1.00	1.00
RDI	1.31	0.81	-2.95	0.73
RMRATE	-0.07	-0.02		
UC			-0.30	-0.02
Time		0.005		0.006

Notes: For definitions of the variables see Appendix 1. Sample period: 1982Q2 to 2000Q1. The order of the unrestricted VAR (lags = 5) has been selected on the basis of the AIC and SW criteria. For Model 1 and 2 the variables included in the cointegrating vector are RHP, RDI, RMRATE. In Model 3 and 4 RMRATE is substituted with the real user cost UC. List of I(0) variables included in the VAR: D8234, DMSY(-1), DMRATE(-1), DMRATE(-2) and seasonal dummies. "*r*" denotes the number of cointegrating vectors. (*) and (**) indicate not rejection of the null hypothesis at 10% and 5% level.

TABLE A.3.2 Johansen Tests for Cointegration for Finland

Model 1

Cointegration with unrestricted intercepts and no trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	47.53	27.42	24.99
r<= 1	r = 2	**16.81****	21.12	19.02
r<= 2	r = 3	9.41	14.88	12.98
r<= 3	r = 4	0.24	8.07	6.50

Cointegration LR Test Based on *Trace* of the Stochastic Matrix

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	73.99	48.88	45.70
r<= 1	r>= 2	**26.46****	31.54	28.78
r<= 2	r>= 3	9.65	17.86	15.75
r<= 3	r = 4	0.24	8.07	6.50

Model 2

Cointegration with unrestricted intercepts and restricted trends in the VAR
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	47.57	31.79	29.13
r<= 1	r = 2	**17.09****	25.42	23.10
r<= 2	r = 3	11.78	19.22	17.18
r<= 3	r = 4	5.29	12.30	10.55

Cointegration LR Test Based on *Trace* of the Stochastic Matrix

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	81.75	63.00	59.16
r<= 1	r>= 2	**34.15****	42.34	39.34
r<= 2	r>= 3	17.08	25.77	23.08
r<= 3	r = 4	5.29	12.39	10.55

TABLE A.3.2 Johansen Tests for Cointegration for Finland (continued)

Model 3

Cointegration with unrestricted intercepts and no trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	39.17	27.42	24.99
r<= 1	r = 2	**18.81****	21.12	19.02
r<= 2	r = 3	7.12	14.88	12.98
r<= 3	r = 4	1.20	8.07	6.50

Cointegration LR Test Based on *Trace* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	66.32	48.88	45.70
r<= 1	r>= 2	**27.15****	31.54	28.78
r<= 2	r>= 3	8.33	17.86	15.75
r<= 3	r = 4	1.20	8.07	6.50

Model 4

Cointegration with unrestricted intercepts and restricted trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	39.80	31.79	29.13
r<= 1	r = 2	**21.90****	25.42	23.10
r<= 2	r = 3	10.35	19.22	17.18
r<= 3	r = 4	1.39	12.39	10.55

Cointegration LR Test Based on *Trace* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	73.46	63.00	59.16
r<= 1	r>= 2	**33.65****	42.34	39.34
r<= 2	r>= 3	11.75	25.77	23.08
r<= 3	r = 4	1.39	12.39	10.55

Estimated Normalised Cointegrating Vectors

	Model 1	Model 2	Model 3	Model 4
RHP	1.00	1.00	1.00	1.00
RDI	1.37	1.50	1.60	1.95
RMRATE	-0.003	-0.004		
UC			0.09	0.09
DEM	1.37	1.49	0.80	0.69
Time		-0.001		-0.003

Notes: For variable definitions see Appendix 1. Sample period: 1981Q2 to 2001Q1. The order of the unrestricted VAR (lags = 5) has been selected on the basis of the AIC and SW criteria. For Model 1 and 2 the variables included in the cointegrating vector are RHP, RDI, RMRATE and DEM. In Model 3 and 4 RMRATE is substituted with UC. List of I(0) variables included in the VAR: D902, D912, DRMSTOCK(-1), DMRATE(-1), DMRATE(-2), seasonal dummies and compounded dummies taking "1" and "-1". "*r*" denotes the number of cointegrating vectors. (*) and (**) indicate not rejection of the null hypothesis at 10% and 5% level.

TABLE A.3.3 Johansen Tests for Cointegration for Sweden

Model 1

Cointegration with unrestricted intercepts and no trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	22.74	21.12	19.02
r<= 1	r = 2	**8.77****	14.88	12.98
r<= 2	r = 3	0.02	8.07	6.50

Cointegration LR Test Based on *Trace* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	31.55	31.54	28.78
r<= 1	r>= 2	**8.80****	17.86	15.75
r<= 2	r = 3	0.02	8.07	6.50

Model 2

Cointegration with unrestricted intercepts and restricted trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	33.39	25.42	23.10
r<= 1	r = 2	**9.68****	19.22	17.18
r<= 2	r = 3	5.09	12.39	10.55

Cointegration LR Test Based on *Trace* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	48.16	42.34	39.34
r<= 1	r>= 2	**14.77****	25.77	23.08
r<= 2	r = 3	5.09	12.39	10.55

TABLE A.3.3 Johansen Tests for Cointegration for Sweden (continued)

Model 3

Cointegration with unrestricted intercepts and no trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	29.35	21.12	19.02
r<= 1	r = 2	**8.37****	14.88	12.98
r<= 2	r = 3	4.27	8.07	6.50

Cointegration LR Test Based on *Trace* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	42.00	31.54	28.78
r<= 1	r>= 2	**12.65****	17.86	15.75
r<= 2	r = 3	4.27	8.07	6.50

Model 4

Cointegration with unrestricted intercepts and restricted trends in the VAR.
Cointegration LR Test Based on *Maximal Eigenvalue* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r = 1	29.36	25.42	23.10
r<= 1	r = 2	**8.84****	19.22	17.18
r<= 2	r = 3	5.26	12.39	10.55

Cointegration LR Test Based on *Trace* of the Stochastic Matrix:

Null	Alternative	Statistic	95% Critical Value	90% Critical Value
r = 0	r>= 1	43.46	42.34	39.34
r<= 1	r>= 2	**14.10****	25.77	23.08
r<= 2	r = 3	5.26	12.39	10.55

Estimated Normalised Cointegrating Vectors

	Model 1	Model 2	Model 3	Model 4
RHP	1.00	1.00	1.00	1.00
RDI	0.62	2.39	0.50	0.60
RLT	-0.09	-0.07		
UC			-0.10	-0.09
Time		-0.008		-0.001

Notes: For variable definitions see Appendix 1. Sample period: 1980Q4 to 2000Q1. The order of the unrestricted VAR (lags = 6) has been selected on the basis of the AIC and SW criteria. For Model 1 and 2 the variables included in the cointegrating vector are RHP, RDI, RLT. In Model 3 and 4 RLT is substituted with UC. List of I(0) variables included in the VAR: D831, D931, DRMSTOCK(-1), DLT(-1), DLT(-2) and seasonal dummies. "r" denotes the number of cointegrating vectors. (*) and (**) indicate not rejection of the null hypothesis at 10% and 5% level.

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CONCLUSIONS

SUMMARY OF FINDINGS, POLICY IMPLICATIONS AND AVENUES FOR FUTURE RESEARCH

1. Introduction

This thesis has addressed some empirical questions related to the transmission mechanism of the monetary policy in a number of industrialised countries over the recent decades. In chapter 1 the focus has been placed upon the effects of unexpected monetary shocks to current account and exchange rate fluctuations in fifteen OECD economies. Chapter 2 has shifted the analysis to the sensitivity of real house prices to short-term interest rate disturbances and their role for household consumption decisions in some European countries. Finally, chapter 3 has attempted to study the main macroeconomic factors driving real house prices in the EU, with a particular emphasis upon the role of real and financial determinants. In the sections that follow, the empirical approaches, the results and the policy implications of the three studies are briefly summarised. At the same time, some of the weaknesses inherent to each are also considered, with the possibility of extensions and avenues for future research.

2. *Results, Implications and Avenues for Future Research*

Chapter 1:
Nominal Shocks and the Current Account: a SVAR Analysis in 15 OECD Countries

The neoclassical intertemporal approach, based on micro-founded optimising models, tends to explain output, prices, current account and other variables dynamics focusing on flexible-price and non-monetary economies, where fiscal and technology shocks are considered to be the main driving factors. Following the publication of the Obstfeld and Rogoff paper in 1995, new insights and testable predictions have appeared in the academic market. In particular, by introducing market imperfections (namely, price stickiness and monopolistic competition), similarly to the traditional aggregative Mundell-Flemming framework, the new models show that monetary policy becomes a potential stabilisation tool and an additional macroeconomic disturbance. Over the recent years, there has been an extensive 'production' of theoretical models relaxing and extending the *Redux* set-up to answer a number of policy-related questions. Although sharing most of the original structure, however, such models have produced different (and, sometimes, diametrically opposed) theoretical predictions of the MTM in open economies. Amongst the most controversial, there are different views upon how unanticipated monetary impulses affect the current account and on the relative importance of the exchange rate channel as equilibrating mechanism.

This chapter has attempted to fill this gap by extending the previous empirical literature, mainly based on the U.S. and the other G-7, and verifying whether current account changes constitute an important channel of the international monetary transmission mechanism in fifteen OECD countries over the period 1979-1998. The implemented econometric approach is based on Structural Vector Autoregression

(SVAR) models, which are the ideal framework for posing the question on the dynamics of the variables of interest following real and nominal structural disturbances. Following the work of Blanchard and Quah (1989) and Clarida and Galì (1994), the latter shocks are identified with a minimum of long-run neutrality restrictions, which are consistent with the findings of previous empirical studies, and imposed or predicted by most macro theoretical models. In particular, two SVAR models are estimated separately for each of the countries under investigation. The first is a simple bivariate specification formed by the first difference of the logarithm of the real exchange rate and the level of the current account to output ratio. In this system, two types of shocks, namely a permanent and a temporary shock, are assumed to drive the two endogenous variables. Their identification is achieved by constraining the latter shock, which consequently is interpreted as a nominal disturbance, to have only temporary effects on the real exchange rate.

In order to test for the robustness of the above model to problems of miss-aggregation of shocks, the two-variable specification is augmented with a measure of relative output. In the new system, the nominal shocks are restricted to have only transitory effects on the level of output and the level of the real exchange rates, whereas the current account dynamics are freely estimated.

The results of the two systems are very similar and unambiguous. In particular, the variance decomposition analysis not only shows a dominant role of nominal shocks for current account fluctuations in most of the fifteen countries, but also clearly indicates that unexpected monetary policy expansions produce temporary current-account improvements, which are accompanied by statistically significant short-run depreciations of the real exchange. Additionally, it is found that the responses of current accounts are positively related to the degree of trade openness of the country, with smaller and relatively more open economies being more affected by nominal disturbances.

These findings have important implications. First, results show that the current account represents an important variable entering the monetary transmission mechanism and, as a result, it would be desirable for the new theoretical frameworks modelling open economies and for policy makers to account for such effects. Moreover, the chapter clearly suggests that expansionary shocks generate temporary (although strongly significant) surpluses of the current account, a result that contrasts with many extensions of the *Redux* model characterised by full pricing-to-market (PTM) behaviour or low degree of substitutability between goods across countries. Temporary deviations from PPP, however, signify that the presence of some degree of PTM are supported by the data. An additional implication is more closely related to the debate over Euroland as an optimal currency area. Exchange rates and external imbalances seem to have represented an important MTM in the EMU economies over the sample period, implying that the latter have not only lost an important re-equilibrating channel to macroeconomic disturbances, but that their current account could be asymmetrically affected by monetary policy actions of the ECB.

The results of the above chapter and the subsequent implications could be extended in two main ways. First, given the quality of the current account statistics, which are based on relatively reliable goods and services data, but less precise income and current transfers, it would be worth testing for alternative current account measures. For instance, the robustness of the estimation results may be checked by using the 'absorption approach' definition of the current account, that is, the difference between national savings and domestic investment. Second, as pointed out in the main text, the parsimony of the estimated systems might over-stress the importance of the identified structural shocks. In this respect, not only alternative identifying restrictions aimed at disentangling more directly nominal disturbances with policy instruments used by the monetary authorities, but also richer models could be implemented. A less parsimonious approach, however, might easily lead to additional noise, over-parameterised specifications and deterioration of the statistical significance of the estimated coefficients. This is the main limitation of the VAR framework, which, although being a powerful tool in macroeconomics, is characterised by a trade-off between the inclusion of additional *a priori* relevant variables in the specifications and accuracy of the estimated parameters.

Chapter 2:
Monetary Policy Shocks and the Role of House Prices across European Countries

Chapter 2 has concentrated on another channel of the monetary transmission mechanism working through policy-induced changes in residential house prices. The last few decades have been characterised by large asset price fluctuations, which are thought to have affected the real activity in many industrialised countries. In this respect, a relevant policy-related question is to assess the role of house price dynamics in the transmission of monetary disturbances to the aggregate demand. While there are many empirical studies on the MTM in European countries evaluating the overall effect of policy changes on the real economy, relatively little effort has been put into the direct estimation of individual channels of the transmission mechanism. Even rarer are studies taking real asset price dynamics directly into account and investigating the quantitative importance of the 'house-price' channel.

Extending the previous VAR empirical literature, this chapter has made a step in this direction. Similarly to the first part of the thesis, alternative VAR models are estimated for a group of European economies, namely Belgium, Finland, Ireland, Italy, the Netherlands, Sweden, Spain and the United Kingdom. These countries represent a very interesting cross section in that they present divergent housing and mortgage markets, which, besides a heterogeneous degree of financial deregulation, are also reflected in differences in the intensity of competition, housing tenure, fiscal treatment of mortgage interest rates, typical features of mortgage contracts, etc. Consequently, it is reasonable to expect a different role of house prices in their respective MTM.

In each system, unanticipated monetary policy shocks are identified through the imposition of contemporaneous restrictions among the endogenous variables and short-term interest rates are directly included as the main policy instrument at disposal of the monetary authorities. From the estimated impulse response functions, it is found that a monetary policy tightening has significant negative effects on real house prices and that the size and the timing of the impulse responses are partly consistent with differences in the housing and credit systems across countries. In particular, the United Kingdom, Sweden and Finland seem to be the countries with a relatively higher sensitivity of residential prices to interest rate shocks, whereas in the remaining economies residential asset prices are relatively more dependent upon real disturbances. The role of these policy-induced house price changes in the MTM to household consumption expenditure is then directly assessed through the implementation of alternative simulation exercises, in which the estimated house-price effect on non-housing consumption expenditure is shut-off. As expected, in the countries where house price are relatively more affected by monetary disturbances, such changes tend to have an amplifying role in the transmission of a monetary shock to household consumer spending. The latter result is not found in most of the remaining economies.

While the above findings might be dependent upon the specific sample period, in which the financial deregulation process of the 1980s and 1990s was much more intense in the Northern European countries, given the current features of the housing and mortgage systems, the results of the chapter indicate that there is an asymmetric role of residential prices on the real economy across the EU members under investigation. However, the recent upward competitive pressure witnessed in the mortgage markets in many Continental countries is a clear signal that real asset prices might be expected to become a significant intermediate variable in the MTM to aggregate demand.

In this respect, given the progressive liberalisation process which is thought to have increasingly enhanced the empirical potency of this channel, an interesting avenue for future research would be to apply a more flexible approach based on time varying coefficients VARs. The latter econometric tool should be better equipped to detect progressive changes in the role of house prices in the MTM. Moreover, in order to improve cross-country comparability, the chapter has applied alike specifications, in which the short-term interest rate is used as measure of monetary policy. In this regard, it might be useful to estimate country-specific systems, which could more closely represent the monetary reaction function and policy instruments of each individual country during the ERM period. As discussed in the previous sections, however, due to relative short sample periods, the quarterly availability of house price indexes and the subsequent low number of observations, less parsimonious specifications are restricted to the rapid loss of degrees of freedom and subsequent inference problems.

Chapter 3:

House price Dynamic in Europe: a Cross-Country Empirical Analysis

This chapter has studied the main macroeconomic forces affecting house price dynamics for most of the EU members, extending the results of chapter 2 as well the previous housing empirical literature, which is mainly confined to the United Kingdom and the Scandinavian countries.

In particular, the first part of this chapter has provided a detailed overview of the EU house price dynamics, the main developments which occurred in the mortgage markets during the financial deregulation process of the 1980s and 1990s, and other institutional features characterising the individual housing systems during the last few decades. From the above analysis, it has been shown that, besides common macroeconomic factors (i.e. the fall of real and nominal interest rates during the pre-EMU period and the increasing financial deregulation process), European housing and mortgage systems are still characterised by a high degree of heterogeneity, which might imply a different sensitivity of house prices to real and financial shocks. Consistent with the presence of mixed mortgage and housing systems across countries, the following sections have extended the existing cross-country housing literature. The latter is generally based on panel estimation techniques in which the parameters associated with the driving factors are imposed to be homogeneous. This assumption, however, seems very restrictive and places clear reservations upon the interpretation and the reliability of the estimated pooled coefficients found in the literature. Similarly, those international house price studies attempting to estimate individual house price equations using annual data over relatively short sample periods have not provided reasonable results, mainly owing to a low number of observations.

Accordingly, the chapter has first formally tested for this heterogeneity assumption by modelling house prices using the annual Bank of International Settlements (BIS) dataset over the period 1970-2001. In particular, two heterogeneous panel estimation techniques have been employed, the first being the traditional Seemingly Unrelated Regression (SUR) approach, the second implementing the Pooled Mean Group (PMG) methods. Both techniques allowed for a direct test of the homogeneity assumption of the coefficients associated with the factors driving residential prices. After systematically rejecting the latter hypothesis, the following sections have shown the results of separate country-specific house price equations, which were estimated using a newly constructed dataset of the ECB containing quarterly information on national house price indexes, real household disposable income, outstanding mortgage stock, typical mortgage interest rates and demographic variables for the period 1980-2001. In order to account for the presence of cointegration amongst the variables included in the specification and to identify both short and long-run determinants of real house prices, an unrestricted error correction mechanism (ECM) approach is implemented for each economy under investigation. To enhance cross-country comparability, similar specifications are chosen. Following a general-to-specific selection approach, the final estimated models show a high degree of heterogeneity across EU countries. In particular, it is found that real disposable income is the most import factor affecting house prices. The corresponding long-run elasticity, however, differs very greatly from country to country. Besides demographic effects, which seem to be influential in only few countries, another important factor particularly relevant to explain long-term house price dynamics are real mortgage interest rates. Consistently with the results of chapter 2, the latter are quantitatively more important in those countries with relatively more competitive and liberalised housing finance markets (namely the U.K., Denmark, Finland, Sweden and the Netherlands). Additionally, in these countries, it turns out to be very important to control for mortgage market deregulation effects, which, not surprisingly, are found to appear very significant over the sample period. In all equations, regardless of the specific national housing finance system, nominal mortgage interest rate changes have been found to have a strong predictive power, indicating that the rapid falls experienced during the pre-EMU period might have played a powerful contribution to the house price booms witnessed in most European countries during the second half of the 1990s.

Overall, although partly affected by a low comparability and quality of housing statistics, the chapter shows that European house prices are driven by similar determinants. The relative importance of the latter, however, is very heterogeneous across countries. In particular, in the economies where the mortgage market is relatively more liberalised and competitive, housing demand is more sensitive to mortgage interest rate movements. This implies that, under inelastic housing supply, the actions of monetary authorities can have significant effects on residential prices. While stronger market integration and growing financial deregulation within the euro-area will probably increase the efficiency of the housing and the competition in the mortgage markets, EU countries are, however, still very heterogeneous and segmented. As a result, housing cycles are asymmetrically driven by real and financial factors, implying the possibility of a diverse impact on aggregate demand and inflation, financial stability and the supply side of the economy.

There are several extensions that could be made to the above analysis. First, due to data availability at the time of writing this dissertation, house price models in which private housing stock and financial wealth effects are directly included could provide more reliable results. Additionally, demand and supply sides of the EU housing markets can be simultaneously modelled to account for the sensitivity of private residential investment to house price fluctuations and the effects of the investment changes on the stock and, eventually, on prices. Besides these extensions, it might be useful to test whether, following the progressive deregulation of the mortgage market, house price have become increasingly linked to the cost of borrowing and how the role of disposable income on residential asset prices has changed over time. The latter could be achieved by implementing a time varying coefficient approach, where the underlying data-generating process is treated as unstable. In this respect, not only would structural changes which occurred during the financial market deregulation, the removal of mortgage market constraints and the implementation of specific housing tax reforms be better modelled, but the use of State Space models should also help to identify progressive changes in the sensitivity of house prices to fundamentals. Finally, an avenue for future research would be to consider the possibility of non-linear adjustment by using an asymmetric error correction model. This is justified by the fact that house prices could respond differently to when prices are below their long-run equilibrium compared to when they are above it.