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THE EMPIRICAL RELEVANCE OF A BASIC STICKY-PRICE INTERTEMPORAL MODEL

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Glasgow, August 2001

Abstract

In this paper, we first outline the monetary version of the sticky price intertemporal model of Obstfeld and Rogoff (1995, 1996), in which monetary shocks unambiguously generate a permanent nominal exchange rate depreciation and a temporary current account surplus. We then empirically investigate these theoretical predictions in two structural VAR systems for 15 OECD countries over the period 1979-1999, using the long-run restriction identification scheme suggested by Clarida and Galì (1994). Our empirical findings support the main predictions of the basic model, as well as suggesting that monetary shocks play an important role in the current account fluctuations. Moreover, we find that more open economies show greater sensitivity of the current account to monetary shocks.

Keywords: Structural VAR, real exchange rate and current account.

JEL Classification: C32, E40, F41, F42

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* I am grateful to Ulrich Woitek for providing the GAUSS codes and help in implementing the econometric procedure. I also thank Jim Malley, Campbell Leith, Jordi Galì, Phillip Lane, seminar participants at the University of Glasgow and discussants at the Scottish Doctoral Programme Methodology Conference 2001, the European Economic Association Conference 2001 (Lousanne) and the MMF Annual Conference 2001 (Belfast) for useful comments and suggestions. The usual disclaimer applies.

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 model. To this regard, a number of studies published in the early 1980s moved towards an "*intertemporal approach*", based on microfounded optimising models, where preferences, technology and capital markets access are directly included.

Most of this literature, however, addresses the intertemporal analysis of the current account determination focusing on flexible-price and non-monetary economies, in which fiscal and technology shocks are the main sources of current account fluctuations (Glick and Rogoff, 1995 and Obstfeld and Rogoff, 1996a for a survey). Only recently, following the partial failure of this class of models in explaining the high volatility of the current account, some studies have introduced a role for monetary shocks as fundamental determinants of external imbalances. The *Redux* model by Obstfeld and Rogoff (1995, 1996) (henceforth OR) is without doubt considered as the precursor of this literature, which introduces market imperfections (namely price stickiness and monopolistic competition) in an intertemporal optimising open economy framework.[1] Their model shows that, under floating exchange rates, monetary shocks unambiguously affect consumption, income and current account.

However, the original assumptions of the OR framework have been relaxed by many authors, who show that the main predictions are not robust to alternative hypothesis on the key parameters (see Lane, 2001 for an excellent survey). As a result, we conduct some empirical evidence on this direction and evaluate the relative importance of monetary shocks in explaining real exchange rates and current account fluctuations in 15 OECD countries by estimating two structural VARs. In order to identify the structural forces driving the variable dynamics over the short--run, we make use of long-run identification restrictions *à la* Blanchard and Quah (1989) and Clarida and Galì (1994).

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[1] For several years, the traditional aggregative Keynesian models with sticky prices and the flexible-price intertemporal neo-classical models have represented the main way of modelling the economy. Both of them, however, present weaknesses. The former ones, although they allow for nominal rigidities, have many shortcomings in what they lack of microfundations for intertemporal choice by ignoring the intertemporal budget constraints, say very little about current accounts and budget deficits, and, in general, do not provide a clear description of the transmission mechanisms of the monetary and fiscal policies. Moreover, they do not allow for welfare and normative analysis which enormously limit the possibility of formulating policy prescriptions. On the other hand, the neo-classical models, while embodying many of these central issues, are still based on the unrealistic hypothesis of flexible prices and perfectly competitive markets.

The aim of this study is twofold. Firstly, as pointed out by Lane (1999), this exercise can be helpful in distinguishing between those class of models predicting a current account surplus in response to a monetary shock (Mundell, 1983 and Obstfeld and Rogoff, 1995, 1996 among others), and those in which a current account neutrality or deficit might occur as in Betts and Devereux (1996, 2000), Lombardo (2001), Kollmann (1997) and Chari, Kehoe and McGrattan (1997), Tille (2001). A second aim is to investigate the role of nominal or monetary shocks in the understanding of the current account determination, which, in the previous empirical analysis, has mainly been focusing on technology and fiscal shocks (Glick and Rogoff, 1995).

The structure of the rest of the paper is organised as follows. In Section 2, we briefly outline the monetary version of the OR model and some of its predictions. Section 3 summarises two main extensions, which, by keeping the basic structure of the *Redux* model, while introducing pricing-to-market and different degrees of elasticity of substitution between home and foreign goods, allow for the possibility of different qualitative implications. Section 4 provides an overview of some recent empirical attempts to test the role of monetary shocks on exchange rate and current account determination. Section 5 outlines the econometric approach we use, while Section 6 comments on the specifications and provides the estimation results. Section 7 concludes.

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

In what follows we outline a textbook version of the *Redux* model in which Obstfeld and Rogoff (1995, 1996) develop a perfect foresight two-country general equilibrium monetary model which combines three fundamental blocks.[2]

1. the first emphasises the <u>*intertemporal decisions by individual agents*</u> where foreign trade and asset exchange open up avenues for transferring resources over time which are not available in a closed economy;
2. the second is based on <u>*monopolistic competition in the goods market*</u>, which plays a central role because it rigorously defends the Keynesian assumption that output is demand-determined in the short-run;

[2] This outline is strongly based on the monetary version model summarised in Chapter 10 of Obstfeld and Rogoff (1996) and the original paper in Obstfeld and Rogoff (1995)

3. the third contemplates the *presence of sticky prices*.

In their model, 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 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 preferences defined over a consumption index, real money balances, and effort expended in production.[3] In particular, the intertemporal utility function of a typical home agent *j* depends positively on consumption and real balances, and negatively on work effort, which is positively related to output. For analytical reasons, OR take into account the specific case, in which all individuals have the same preferences defined by

$$(2.1) \qquad U_t = \sum_{s=t}^{\infty} \beta^{s-t} \left[\log C_s + \frac{\chi}{1-\varepsilon} \left(\frac{M_s}{P_s} \right)^{1-\varepsilon} - \frac{\kappa}{2} y_s(z)^2 \right],$$

where $0 < \beta < 1$ is the subjective rate of discount, $\varepsilon > 0$ is a parameter inversely related to the elasticity of money demand, the elasticity of intertemporal substitution is equal to 1 and the elasticity of disutility from output equal to 2.[4]

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

$$(2.2) \qquad C^j = \left[\int_0^1 c^j(z)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta}{\theta-1}},$$

where the elasticity of substitution between varieties $\theta > 1$.[5] The foreign consumption C^* is defined analogously.

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[3] The assumption of time-separable utility function is preferable for a number of reasons. In particular, Obstfeld and Rogoff (1996) point out that an intertemporal non-separable utility function would yield few concrete and testable predictions, and that empirical research has not been able to provide a superior non-separable alternative.

[4] It is easy to show that a rise in home productivity can be captured by a fall in *k*.

[5] As pointed out by Obstfeld and Rogoff (1995) this assumption is necessary because the parameter θ is also the price elasticity of demand faced by each monopolist. They have to impose this restriction, because since

The price deflator for nominal money balances is the consumption-based money price index corresponding to (2.2).[6] Letting *p(z)* be the home-currency price of good *z*, it is easy to show that the money price level in the home country is:

$$(2.3) \qquad P = \left[\int_0^1 p(z)^{1-\theta}\, dz\right]^{\frac{1}{1-\theta}}$$

Being *p*(z)* the foreign-currency price of good *z*, the foreign price index *P** can be written similarly.

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, then the *law of one price* holds for every good,[7] so that

$$(2.4) \qquad p(z) = Ep^*(z)$$

Since both countries' residents have the same preferences, equation (2.2) implies that home and foreign consumer price indexes are related by *purchasing power parity* (PPP)

$$(2.5) \qquad P = EP^*$$

There is *no capital or investment*, but this is not an endowment economy because labour supply is elastic. This means that period *t* output of good *z*, $y_t(z)$, is chosen in a manner that depends on the marginal revenue of higher production, the disutility of effort, and the marginal utility of consumption, that is, the output is endogenous.

There is an *integrated world capital market* in which both countries can borrow and lend. The only asset they trade is a real bond, denominated in the composite consumption good. Let 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*.

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marginal revenue is negative when elasticity of demand is less than 1, θ > 1 ensures an interior equilibrium with a positive level of output. However, as pointed out by Lombardo (2001), the elasticity of substitution between domestic and imported goods in OR model is "unnecessary" constrained to be bigger than one. He relaxes this assumption providing interesting insights. Tille (2001) further extends the basic model allowing for different elasticity of substitution across and within countries. See the next Section for a brief outline of these extensions.

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

[7] In this model the law of one price always holds. However, several authors (Engel, 1998 – Rogoff, 1996) have documented that international deviations in tradable prices are responsible for a large proportion of real exchange fluctuations. Following this empirical evidence, Betts and Devereux (1996, 1998, 2000) have introduced market segmentation into the basic model in the form of "pricing-to-market" (PTM). As above, this extension produces theoretical implications in the international transmission of monetary (as well fiscal) policy which differ from the OR model. Similarly, Chari *et al.* (1997) develop a sticky price model with price discriminating monopolists, which produces deviations from the law of one price. See the next Section for a brief discussion.

Individual z's period budget constraint written in nominal terms is:

(2.6) $$P_t F_t + M_t = P_t(1+r_{t-1})F_{t-1} + M_{t-1} + p_t(z)y_t(z) - P_t C_t - P_t T_t$$

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 (which can be negative in the event of money transfers).

Since *Ricardian equivalence* holds in this model, they assume that the government runs a balanced budget each period. Therefore, all seignorage revenues are rebated to the public in the form of transfers:

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

The same is true for the foreign country.

Given the constant-elasticity-of-substitution consumption index, eq. (2.2), it is easy to find the home individual j's demand for good z in period t and the correspondent demand of a foreign individual.

Integrating demand for good z across all agents (that is, taking a population weighted average of home and foreign demands), and making use of eqs. (2.2) and (2.4), which implies that $p(z)/P = p^*(z)/P^*$ for any good z, we can determine the constant-elasticity-of-substitution *total world demand* for good z

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

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

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

where they imposed symmetry on the identical agents within each country in order to simplify notation.

From the first order condition of the utility maximisation, the following functions are derived:

(2.10) $$C_{t+1} = b(1+r_t)C_t$$

(2.11) $$\frac{M_t}{P_t} = \left[cC_t\left(\frac{1+i_t}{i_t}\right)\right]^{1/e} \quad ,$$

where we use the *Fischer parity equation* defined as $(1+i_t)=\frac{P_{t+1}}{P_t}(1+r_t)$, and

$$(2.12) \qquad y_t(z)^{\frac{q-1}{q}} = \frac{q-1}{qk}(C_t^w)^{1/q}\frac{1}{C_t}$$

which represent respectively, the *standard Euler equation*, the *money demand and labour supply* equation. The same is true for the foreign country counterpart equations.

In order to analyse the effects of a monetary shock, they carry out a log-linear approximation around a flexible-price steady-state equilibrium (where all the exogenous variables are constant and where the initial stock of net foreign assets is 0, i.e. $\overline{F}_0=\overline{F}_0^*=0$) and solve for differences between home and foreign log-linearized expressions.[8]

In the short-run, nominal producers 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. In this sense, we can interpret the possible sources of stickiness with menu costs of price adjustment. With sticky nominal prices, output becomes demand determined for small enough shocks, because the monopolist always prices above marginal cost, and it is profitable to meet unexpected demand shocks at the pre-set price. Since prices are fixed for one period, Obstfeld and Rogoff distinguish between the impact (first-period) effect of a shock and its long-run (second-period) steady-state effect.

In considering the effects of an *unanticipated permanent rise in the relative home money supply* (that is, $\hat{M}_t-\hat{M}_t^*=\hat{M}_{t+1}-\hat{M}_{t+1}^*$) the main findings and dynamics of the model 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, followed by a decrease of the domestic relative prices. As a consequence, and under the assumption of monopolistic competition, domestic output temporarily expands. With consumption which is based on permanent income, 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

[8] They implement this linearization by expressing the model in terms of deviations from the baseline steady-state path. By denoting percentage changes from the baseline by hats, for any variable, $\hat{X}_t \equiv dX_t/\overline{X}_0$, where $\overline{X}_0$ is the initial steady-state value.

country. Domestic consumption, therefore, rises although the increase in output lasts for only one period.

Although a full analytical description of the relevant equations would make the analysis more thorough, the main findings, which are relevant for the theoretical relationship between monetary policy, exchange rate and current account, can be summarised in the following expressions.

(2.13) $$\hat{E}_t = \hat{E}_{t+1} = (\hat{M}_t - \hat{M}_t^*) - \frac{1}{e}(\hat{C}_t - \hat{C}_t^*)$$

which shows that the exchange rate jumps immediately to its new long-run equilibrium following a permanent relative money shock. The assumption behind this result is that if consumption differentials and money differentials are both expected to be constant, then agents must expect a constant exchange rate as well.

Another important equation, which derives directly from the budget constraint, is

(2.14) $$\frac{dF_t}{\overline{C}_0^w}\left(\frac{1}{1-n}\right) = (\theta - 1)\hat{E}_t - \left(\hat{C}_t - \hat{C}_t^*\right)$$

which tells us 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, but in addition we also have a consumption effect which tells us that the larger the increase of relative consumption, the smaller the wealth transfer.[9]

Eq. (2.14) can be rewritten as follows

(2.15) $$\frac{dF_t}{\overline{C}_0^w} \equiv Z_E = (1-n)(\theta - 1)\hat{E}_t - (1-n)\left(\hat{C}_t - \hat{C}_t^*\right),$$

where we can see that the smaller the home country (i.e. the smaller *n*), and the larger the elasticity of substitution between varieties, the larger the impact on the home current account.

The second expression derived from the model is given by

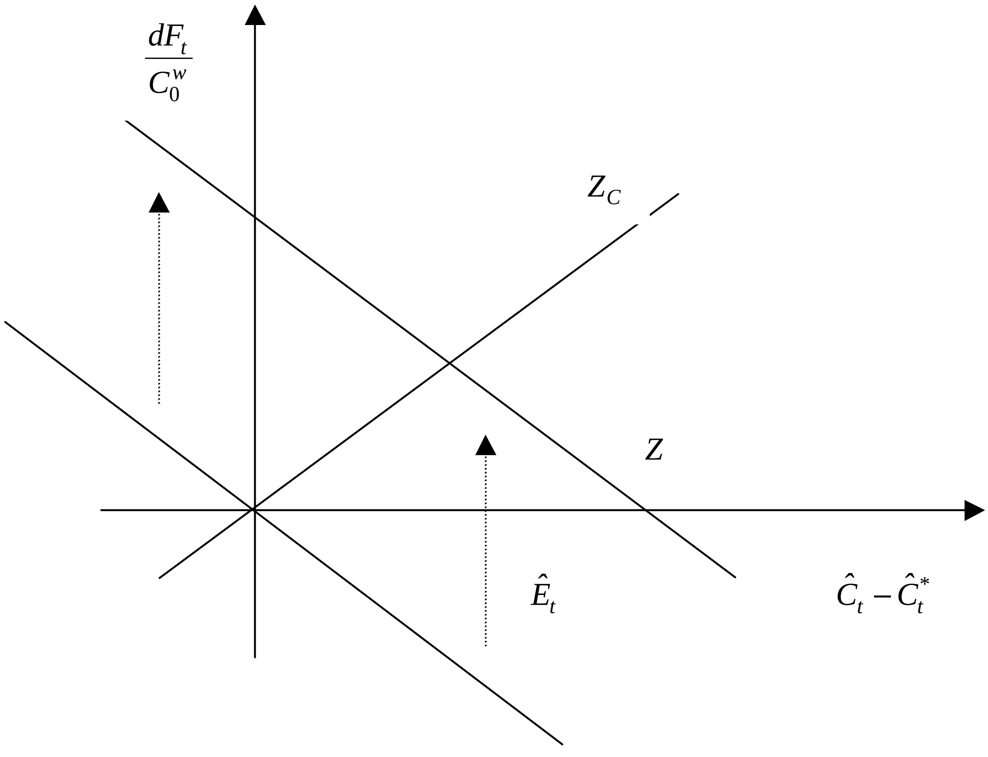
(2.16) $$\frac{dF_t}{\overline{C}_0^w}\left(\frac{1}{1-n}\right) = \left(\frac{2\theta}{\overline{r}(\theta - 1)}\right)\left(\hat{C}_t - \hat{C}_t^*\right),$$

which shows that the consumption change is positively related to the current account surplus through the permanent interest income which home individuals earn from the wealth transfer. Moreover, it is easy to see that the increase in consumption is less than the amount $\bar{r}dF$ (since θ >1), for the fact that the higher wealth leads to some increase in leisure and, consequently, in some reduction of production. Similarly to the expression Z_E, we can rewrite eq. (2.16) as

$$\frac{dF_t}{\overline{C}_0^w} \equiv Z_C = \left(\frac{2\theta(1-n)}{\bar{r}(\theta-1)} \right)\left(\hat{C}_t - \hat{C}_t^*\right) \tag{2.17}$$

We provide a representation of the two lines (namely, Z_E and Z_C) in Figure 1. We can see how their intersection implies that an exchange rate depreciation improves the current account over the short run.

FIGURE 1 Effect of an unexpected home money shock on the current account.



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

3. *Two Extensions of the Basic Model*

As originally Obstfeld and Rogoff pointed out, their framework can be improved with a number of extensions.[10] 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. In what follows, we provide a brief overview of two of the most recent contributions, which, while keeping the basic structure of the model, develop two very interesting extensions by introducing PTM and allowing for different elasticities of substitution across and within countries. It turns out that these assumptions strongly affect the international transmission of monetary shocks and provide different theoretical predictions and welfare implications.

Betts and Devereux (2000) develop an extension of the OR model by allowing short-run departures by 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 s provides a measure of "pricing-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.[11] Betts and Devereux (2000) show that the nominal price stickiness and the presence of PTM increase the volatility of the exchange rate and strongly affect the international monetary transmission mechanisms. In particular, they show that high degree of PTM reduces the traditional "expenditure switching" effects of exchange rate depreciation, which, as a result, has little effect on the relative price of imported goods facing domestic consumer and on the correspondent shift in world demand. Moreover, short-run deviations from PPP tend to generate low (with respect to models without PTM) comovements of consumption and high positive comovements of output across countries. The degree of PTM is clearly central to their results. In particular, they

[10] 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.

[11] Following the lack of empirical support for the law of one price 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. Engel (1993) and Engel and Rogers (1996) provide some

consider the two opposite cases of $s \to 0$ and $s \to 1$. In the former (i.e. the law of one price is maintained continuously and PPP holds) for both countries, a devaluation unambiguously improves the current account, consistently with the traditional OR model, and generates permanent positive effects in the home consumption relative to foreign consumption. When $s \to 1$ and full-PTM occurs, a monetary disturbances generate exchange rate devaluation, but no current account effects. In general, the higher the degree of PTM, a monetary shock implies a greater exchange rate volatility, a smaller current account improvement, and reduced permanent effects on relative consumption.[12]

Another simple, but crucial, extension has been developed by Tille (2001), who allows for the elasticity of substitution across and within countries to differ and generalises the baseline *Redux* model where the two parameters are equal and constrained to be bigger than one (i.e. $q > 1$). Under the assumption of law-of-one-price and PPP, they distinguish 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 MLR (namely Marshall-Lerner-Robinson condition), in that, when goods produced in different countries are close substitutes, an exchange rate depreciation generates a current account surplus, a permanent rise of home consumption relative to foreign consumption. When the goods produced in the two countries are poor substitutes and the elasticity is less than unity (i.e. NON-MLR), the current account will be in deficit in the short run, determining a permanent fall in relative consumption. Similarly, 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 basic Obstfeld and Rogoff's *Redux* model. However, the size and the sign of the effects of monetary shocks are strongly affected by the magnitude of several key coefficients and the structural assumptions. As pointed out by Lane (2001), the above literature has been mainly

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empirical evidence and show that deviations from the law-of-one-price across international borders are greater than can be due to geographical distance and transportation costs.

[12] Similarly to Obstfeld and Rogoff (1995), Betts and Devereux (2000) also focus on the implications of government spending shocks. They show that both temporary and permanent fiscal shocks generate short-term real depreciation and fall of relative consumption. The current account effects crucially depend on the degree of

focusing 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 mainly based on calibration methods and others on the estimation of the key parameters of the model, a number of recent papers have addressed this deficiency by estimating impulse-response functions generated by VAR econometric techniques, where different identification solutions have been used. In what follows, we provide a brief overview of the VAR empirical evidence

4. Empirical Review

The two most influential papers on which the following literature has built are Eichenbaum and Evans (1995) (EE) and Clarida and Galì (1994) (CG). Although they use different identification solutions, they both investigate the effects of shocks to monetary policy on exchange rate in a manner that is qualitatively consistent with the above sticky-price models. In particular, EE use U.S. monthly data covering the sample period 1974:1-1990:5 to estimate alternative VAR models, which share common variables (namely, industrial production, CPI, short-term interest rate differential and exchange rates), but differ for the use of three measures of shocks to U.S. monetary policy.[13] Their standard Cholesky decomposition identification assumes that all shocks are orthogonalised and is such that the monetary variables are ordered prior to the industrial production and the consumer price level and after the interest rate and exchange rate variables. This corresponds to the assumption that the monetary authority sets its policy instrument with current values of the first two variables in mind, while these do not respond contemporaneously to movements of the monetary shock. They find strong evidence that contractionary policy shocks lead to significant and persistent appreciations in exchange rate, both nominal and real, and conclude by pointing out that monetary shocks contributed significantly to the overall variability of U.S. exchange rates in the post-Bretton Woods era.

PTM, the duration of the shock and the magnitude of the elasticity of labour supply. See Betts and Devereux (2000), page 233-35 for a discussion.

[13] They use orthogonalised shocks to the federal funds rate, orthogonalised shocks to the ratio of non-borrowed to total reserves and changes in an index they construct as monetary policy proxy.

Betts and Devereux (1999), by using U.S. data with respect to the remaining G-7 countries, modify the EE's specification, and also consider a short-run recursive identification scheme. They find that positive innovations to monetary policy cause a sharp and persistent depreciations of the real and nominal exchange rates.[14]

Lane (1999) extends the EE system to include a trade balance in order to employ data at a monthly frequency over 1744:1-1996:12 for the U.S. with respect to other G-7 countries. In particular, he estimates a six-variable VAR system, imposing a set of exclusion restrictions on the contemporaneous relationship between the variables. The trade balance is ordered last in the system to allow for contemporaneous effects of all shocks on it. However, his results are robust to alternative orderings and two measures of the US monetary policy instrument (i.e. the Federal Funds rate and the level of non-borrowed reserves). 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 after 43-48 months, according to the monetary policy instrument used.

Another part of the VAR empirical literature has focused on the use of long-run identification scheme *á la* Blanchard and Quah (1989).[15] An influential contribution was made by Clarida and Galì (1994) who investigate empirically and attempt to identify the sources of real exchange rate movements after the collapse of Bretton Woods for the U.S. *vis-à-vis* with the U.K., Germany, Japan and Canada, respectively. They make use of a structural VAR system on output, the real exchange rate and inflation, whose long-run identification restrictions are consistent with a stochastic version of the Obstfeld (1985) open macro model. They assume the presence of three structural shocks to supply, demand, and money and impose restrictions in such a manner that money shocks are expected to have no long-run impact on either output or the real exchange rate, and the demand shock to be long-run neutral for output.[16] They find that for two of the four countries (namely, Japan and Germany), the structural VAR estimates suggest that nominal (money) shocks explain a substantial amount of the variance in dollar-DM and dollar-yen real exchange rates (41% and 35% of the variance, respectively). For the U.K and Canada, variance decomposition

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[14] In Betts and Devereux (1996), they study a Wold decomposition *a la* Eichenbaum and Evans and look at 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.

[15] Se the next Section for an outline of this identification scheme.

[16] See the following Section for a description of their econometric approach.

results are less supportive of the monetary policy shocks relevance. However, consistent with EE, for all countries nominal shocks lead to short-run real depreciation, a rise in U.S. output and a jump in U.S. inflation.

A number of recent papers have made use of the CG's VAR identification scheme, by focusing on the current account rather that the real exchange rate, in line with the qualitative predictions of some of the sticky-price intertemporal model. Lane (1999) estimates a three-variable system where he takes the ratio of the U.S. and rest-of-the-world (proxied by the non-US G-7 countries) output, the US's current account to GDP ratio and the ratio of the U.S. to RoW consumer price levels. He assumes that his system is driven by a sequence of three orthogonal structural shocks, which he labels supply, absorption and monetary shocks, respectively. His long-run identification assumptions are such that the money shock has no long-run effects on the relative output and the current account, and the absorption shock is long-term neutral on the relative output. The estimated IRFs of the current account to a positive one standard deviation monetary shock show 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.

Similarly to Lane (1999), 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, and the monetary shock not to affect output contemporaneously. Estimating this system over 1974:1-1997:4 for the G7 countries, she shows that the current account response to the monetary shock differs across countries. In particular, in the case of UK, Italy, France and Canada, the current account goes into surplus following a negative monetary shock. For the US, Germany and Japan the reverse holds.

Prasad (1999) 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 particular, in accordance with the implications of a modified version of Clarida and Galì (1994) theoretical 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 G-7 countries. In particular, he shows that positive monetary shocks determine significant trade surplus and generate positive correlations between output and trade balance.

Similarly to above, Lee and Chinn (1998) estimate an even more parsimonious two-variable VAR model containing the real exchange rate and the ratio of the current account to GDP for the G-7 countries, and impose a long-run neutrality restriction of a temporary (monetary) shock on the real exchange rate. Their results are supportive of the real exchange rate depreciation and current account surplus following a positive nominal shock.

Many of the above studies focus on transmission mechanism of the monetary shocks to the nominal and/or real exchange rate and current account/trade imbalances mainly in the U.S. and the remaining G7 countries. However, none of them has been applied to a wider set of countries. In what follows, we try to fill this gap by studying the qualitative and quantitative role of monetary shocks in the current account dynamics in a number of additional small open economies. Several studies have shown that for non-US countries, where US and other foreign financial and macroeconomic conditions are strictly under observation by the respective central banks, it is necessary to modify the identification allowing for feedback effects as well as augmenting the specification with a number of key endogenous variables in the system (Kim, 1999, 2000; Kim and Roubini, 1999 among others). To this regard, the VAR identification based on contemporaneous restrictions is not the most appropriate and simple method to apply. On the basis of such considerations, we apply the identification scheme suggested by Blanchard and Quah (1989) and Clarida and Galì (1994) and, as a result, keep our specifications as parsimonious as possible, contemporaneously testing for the main driving forces of current account dynamics. Before motivating our specifications and providing the correspondent empirical results, we outline our econometric approach.

5. *The Econometric Approach*

The empirical strategy we implement builds on earlier work by Blanchard and Quah (1989) and Clarida and Galì (1994), who propose an identification scheme for VARs, in which they impose long-run restrictions on the behaviour of the variables in the system.

In outlining their empirical strategy, we consider a trivariate system, in which the variables are assumed to be stationary. Letting $x_t \equiv [y_t, z_t, q_t]'$ denote the *(3x1)* vector of the system's 3 variables and $e_t \equiv [s_t, d_t, v_t]'$ denote the *(3x1)* vector of the system's 3 structural disturbances,[17] we assume that x_t is a covariance stationary vector process and generated by the following *structural* moving average (MA) model:

(5.1) $x_t = C(L)e_t = C_0 e_t + C_1 e_{t-1} + C_2 e_{t-2} + \ldots$

where C_0 is the *(3x3)* matrix of the contemporaneous structural relationship among y_t, z_t and q_t, and where we assume that the *structural* disturbances e_t are mutually orthogonal and have unit variance, implying that $Ee_t e_t' = I$.

The *reduced-form* MA representation for x_t, which is directly estimates, is given by

(5.2) $x_t = R(L)u_t = u_t + R_1 u_{t-1} + R_2 u_{t-2} + \ldots$

where u_t is a *(3x1)* vector reduced-form disturbance.

Assuming that there exists a non-singular matrix S such that

(5.3) $u_t = Se_t$

and comparing (5.1) and (5.2), we can easily see that

(5.3a) $C(L) = R(L)S$ i.e. $C_0 = S, C_1 = R_1 S, C_2 = R_2 S \ldots$

This implies that (5.3) can be also be written as

(5.4) $u_t = C_0 e_t$.

OLS can be used to obtain consistent estimates of the parameters in (5.2) as well as an estimate of the symmetric variance-covariance matrix of the *reduced-form* disturbances u_t:

(5.5) $Eu_t u_t' = \Sigma$

Therefore, from (5.3), (5.4) and (5.5), and the assumption of mutually orthogonal structural shocks, together with the normalisation condition above, we get

(5.6) $\Sigma = C_0 C_0' = SS'$.

As it is well known in the literature, the system (5.6) provides 9 equations in only 6 unknowns, that is, the 3 variances and the 3 covariances that define Σ. Just-identification of

[17] It is important to point out that, in contrast with the Sims-Bernanke procedure, this identification scheme does not directly associate the three structural shocks with the three sequences {y(t)}, {z(t)} and {q(t)}. In particular, we can think of the latter ones as the endogenous variables, and the three shocks as the exogenous variables of the system.

model (5.2) and, therefore, estimation of the matrix C_0 and the structural innovations e_t require 3 additional restrictions, which will be given consistently with OR *Redux* model.[18]

In particular, letting $C(1) \equiv C_0 + C_1 + C_2 + ...$ denote the matrix of long-run coefficients such that

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

CG assume three long-run neutrality conditions such that $C(1)$ is restricted to be lower triangular. This means that the structural shocks d_t and v_t do not affect the variable y_t in the long run, implying that

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

Similarly, the assumption that structural shocks v_t do not influence the second endogenous variable z_t in the long run requires that:

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

In what follows, we show that (5.8) and (5.9) are sufficient to identify the structural matrix C_0, to recover the structural-system dynamics defined by $C_1, C_2, ...$, as well as the structural shocks e_t.

From (5.3a) it easy to see that $R_0 = I, R_1 = C_1 C_0^{-1}, R_2 = C_2 C_0^{-1}$, and so on. Therefore, the reduced-form MA model (5.2), that we directly estimate, can be rewritten as:

(5.10) $$x_t = R_0 u_t + R_1 u_{t-1} + R_2 u_{t-2} + ...$$

where we can note that

(5.11) $$R(1) \equiv R_0 + R_1 + R_2 + .. = C(1) C_0^{-1} .$$

Provided the above assumptions, CG define the matrix:

(5.12) $$R(1) \Sigma R(1)'$$

which can be computed from the estimation of the variance-covariance matrix of the reduced-form disturbances Σ, and the reduced-form long-run coefficients $R(1)$, which we get from (5.10). Using the definition of $R(1)$ and equation (5.6) to substitute for Σ, it is possible to write:

[18] We will discuss the actual interpretation of the structural disturbances in the next Section.

(5.13) $R(1)\Sigma R(1)^{'} = C(1)C(1)^{'}$.

Letting H denote the lower triangular Cholesky decomposition of (5.13) and knowing that this matrix is unique, we have that $HH^{'} = R(1)\Sigma R(1)^{'} = C(1)C(1)^{'}$ implying:

(5.14) $C(1) = H$

since we have imposed the structural long-run coefficient $C(1)$ to be also lower triangular. Finally, given (5.14) and (5.11), we can obtain

(5.15) $C_0 = R(1)^{-1} H$

which allows us to compute the structural dynamics defined by $C_1, C_2, \ldots$ as well as the structural shocks e_t.

6. Specification and Estimation Results

The two specifications we use closely follow the bivariate VAR system of Lee and Chinn (1998) and the trivariate model in Prasad (1999), in which they impose a long run neutrality condition of monetary or nominal shocks 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.[19] While this is true in many theoretical extensions of the *Redux* structure, their neutrality restriction is trivially consistent with all monetary and Keynesian aggregative models, as well as the theoretical frameworks of Section 2 and 3, provided an appropriate definition (or interpretation) of the real exchange rate.[20]

An important advantage of our identification relies on the fact that we do not impose any long-run neutrality restriction on the current account. In the intertemporal approach, any real

[19] There is a large empirical literature emphasising the long-term implications of cumulated current account imbalances (and net foreign assets) on the real exchange rate (see a recent work from Lane and Milesi-Ferretti, 2000). In this paper, we mainly focus on the short-run dynamics of real exchange rates and current account, and, consequently, abstract from these long run effects.

[20] In the basic OR model, free trade implies that the LOOP holds for individual goods. Moreover, given that consumers have identical preferences across goods and identical CES utility functions, PPP holds in terms of consumer price indices. 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 large non-tradable component, we take the CPI-deflated exchange rate as a proxy for the real exchange rate.

and nominal shock typically are assumed to have no long run effect on this variable. As a result, we allow the system to freely determine its dynamics, which, therefore, can be considered over-identifying restrictions useful to interpret and disentangle the structural shocks. Moreover, the unit root tests strongly reject the non-stationarity hypothesis of the current account to GDP ratio. This result, which is consistent with intertemporal models, clearly suggest that making the assumption that real shocks are the source of a unit root in this variable and that the nominal shock has a long-run neutrality effect on the current account might not be advisable.

In what follows, we show the results of two estimated SVARs based on the long-run restriction identification scheme outlined in the above Section. Data description and unit root tests are in Appendix 1.

Estimation of the bivariate VAR

The first specification is an extension of Lee and Chinn (1998) bivariate VAR $x_t = \left[\Delta REER_t ,\ CA_t / Y_t\right]$ to 15 OECD countries,[21] which we use as a benchmark specification for testing the robustness of the results in the trivariate system of the next section. In this model, we assume the existence of two types of structural shocks generating current account and real exchange rate dynamics: a *permanent* shock and a *temporary* shock. In this framework, we can think of *permanent* shocks as *productivity* disturbances, and temporary shocks as *monetary* shocks.[22] On the basis of this interpretation, we constrain the latter disturbance so that it has no long-run effect on the real exchange rate variable. In doing so, we force the PPP to hold in the long run, although we allow for short-term deviations as in the PTM extension.[23] The advantage of such a model is on its simplicity, as well as on the fact that it is capable of identifying the impact of the two shocks on the current account and exchange rates, without constraining the short-run dynamics or the contemporaneous exclusion of any structural shocks from any equation.

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[21] Namely, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Portugal, Spain, Sweden, the U.K. and the U.S.

[22] This specification does not allow us to distinguish between money demand and money supply shocks. In order to do this, we would have needed to include monetary variables and imposed additional long-run (and/or short-run) restrictions.

[23] This restriction is consistent with the PPP assumption of the basic OR model and the generalisation by Tillie (2001) and Lombardo (2001), as well as with Betts and Devereux (2000) where short-run fluctuations of the real exchange rate are allowed in the short run. In doing so, we can observe the dynamics as well as the statistical significance of short-run deviations from PPP.

This system is estimated independently for each of the 15 OECD countries, by taking the first difference of the REER series and the level of the ratio CA/Y, in accordance with stationarity properties of the series.[24] Similarly to Lee and Chinn (1998), we use two lags, which is a fair balance between the lag lengths chosen 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 selects 2-3 or, in most of the cases,

more lags. However, with the exception of Denmark where we use 4 lags, our results are qualitatively robust to alternative lag lengths. To save space, and given that results are very similar across countries, Figures 2-5 show the impulse response functions (IRFs) of the *level* of the real exchange rate and the ratio of the current account to output in response to the two different types of shocks for four countries: Finland, France, Sweden, the U.S. [25] The dotted lines are 95% confidence interval bands obtained with the bootstrap-after-bootstrap procedure described in Kilian (1998), which accounts for the bias of small-sample distribution of the impulse response parameters.

The results, which for the G-7 are obviously similar to Lee and Chinn (1998), show that a positive one standard deviation temporary shock generates a statistically significant short-run improvement of the current account. 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. Differently from Lee and Chinn (1998), however, these point estimates are not statistically significant. This result might derive from the different method used in estimating the confidence bands. The U.K. and Belgium provide some anomalous results in that, despite the presence of an external balance surplus to a monetary shock over the short-run, the level of the exchange rate appreciates, although not statitically significant.

Although permanent shocks are not the object of this analysis, it is worth pointing out that, in all countries, these shocks improve the current account (with the exception of Belgium and the U.K.). However, these results might be due to the parsimony of the model, which does not include other key variables of the transmission mechanism or distinguish between

[24] Other studies find that the current account to output ratio has a unit root (Lane, 1999; Cavallari, 2001). Although in our sample period we cannot reject its stationarity (see Appendix), we have re-estimated the system with both variables in first-difference. Results are qualitatively similar, but generate permanent effects of the identified shocks on the current account.

other real shocks. As pointed out by Lee and Chinn (1998), however, the fact that the two identified shocks generate opposite correlations between the current account and the exchange rate suggests the centrality of the nature of the shock in disentangling this relationship. The estimated IRFs, however, are in line with the empirical studies outlined in Section 4 and provide some qualitative support on the interpretation of temporary shocks as monetary ones.

Finally, Table 4 shows the variance decompositions to illustrate the quantitative contribution of the two shocks to exchange rate and current account volatility for the four representative countries.[26] According to our estimates, for most of the 15 OECD countries the nominal shock plays a greater role in explaining the variation of the current account in all the 24 period horizons (i.e. over 70-80%). The only exceptions are France and, in particular, the U.S., where the role of monetary disturbances for external imbalances is minimal (around 20% on average). Contemporaneously, real shocks dominate the volatility of the real exchange rates in all countries.

Estimation of the trivariate VAR system

In order to verify the robustness of the above results, we augment the above specification with a key endogenous variable (i.e. the relative output) and estimate a three-variable system $x_t = \left[\Delta(Y_t / Y_t^*), \Delta REER_t,\ CA_t / Y_t\right]$ for the 15-OECD countries. We chose this specification for a number of reasons. Firstly, the simple bivariate system might be subject to problems of misaggregation of shocks. This might be suggested by the fact that the "identified" monetary shock accounts for most of the variance of the current account. Our specification allows us to identify, and possibly disentangle, an additional *real* shock which can have only *temporary* effects on the relative output. As in Clarida and Galì (1994), Prasad (1999) and Lane (1999), we can interpret this innovation as a non-monetary demand or absorption shock. In doing so, we disentangle a further leading force affecting the endogenous variables and "improve" the correct identification of the monetary shock, which is assumed to have no long run effect on the relative output and the real exchange rate.

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[25] The above four countries are a mix of small and large economies, as well as a combination of countries which over the sample period operated in a flexible exchange rate and semi-fixed exchange rate regime. The remaining countries IRFs and results are available from the author upon request.

[26] Results are available from the author upon request.

Similarly to the bivariate VAR, we choose a lag length which represents a good balance between the AIC and the SIC.[27] In order to allow for a comparison with previuous empirical studies on the current accout determination, Figures 6-12 report the impulse responses of the current account to GDP ratio in response to a positive one standard deviation of the three structural shocks in the G-7 countries.[28]. The dotted lines are 95% confidence interval bands obtained with the bootstrap-after-bootstrap procedure described in Kilian (1998), which accounts for the bias of small-sample distribution of the impulse response parameters. Given the focus of our analysis, we do not present the impulse responses for the relative output and the real exchange rate, which, however, are quite reasonable.[29]

In particular, we find that in all G7 countries the current account is positively and temporarily affected by the positive monetary shock. This improvement is statistically significant in all countries and confirms the results obtained in the bivariate system. It also supports the basic prediction of the basic *Redux* model and the correspondent extensions characterised by low degree of PTM behaviour and high elasticity of substitution between domestic and imported goods. The responses of the current account to the supply and demand shocks vary across countries, although they appear to be statistically insignificant in most of the cases. In particular, positive supply shocks generate a decreasing improvement of the current account ratio (except for Canada, Italy and the U.K.). Similarly, demand shocks improve the current account in the U.S., Canada, Japan and the U.K. and decrease it in Germany, France and Italy, although the effect is small and not statistically significant.[30]

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[27] Accordingly, we chose a lag length of four for Austria, Italy, Japan, Spain, the U.K and the U.S., six for Canada, two for Belgium, France, Germany, Portugal and Sweden, and one for Denmark, Finland and the Netherlands. Our estimates, however, are qualitatively similar regardless of the lag length used.

[28] The remaining OECD countries show similar results, which are available from the author upon request.

[29] The identified permanent shock, which could be interpreted as a relative supply or technology shocks, permanently and significantly increases output. It also generates a persistent and significant (except for the U.K., Italy and Germany) appreciation of the real exchange rate. This result is also found in our bivariate VAR and other empirical studies, which use similar methods (Prasad, 1999 and Clarida and Galì, 1994) or contemporaneous restriction identification methods (Muscatelli and Spinelli, 1999). The non-monetary demand shock causes a permanent and statistically significant appreciation of the exchange rate and tend to temporary, but not significantly, increase the relative output in most countries. Finally, again with the only exception of the U.K., the nominal shocks lead to temporary and statistically insignificant depreciation of the real exchange rate. The only puzzling result is given by the temporary decrease of the relative output following the monetary shock for some of the countries under investigation. The correspondent IRFs, however, are not statistically significant.

[30] These results do not help us to provide a clear-cut interpretation of this shock. However, in the *Redux* model, Obstfeld and Rogoff (1996) analyse the effects of permanent positive home government spending shocks which are qualitatively consistent with the estimated IRFs.

Table 5 shows the variance decomposition of the current account. We can easily notice the new system confirms the quantitative importance of the monetary shock. In fact, apart from France, Germany and the U.S. where the supply shock is dominant,[31] nominal disturbances still account for more than 50 per cent in Canada, Italy, Japan and the U.K. over the 24 period horizon. The results (not shown) for the remaining 15 OECD economies are even more supportive, in that the monetary shock dominate the variance decomposition of the current account (40-80 %). Differently from the bivariate specification, the demand shocks dominate the forecast error variance decomposition of the real exchange rate. The only exceptions are Japan and the U.S., where supply and nominal shock, respectively, provide the major contribution. Not surprisingly, these final results suggest that the relative importance of these shocks might be sensitive to different specifications. However, both systems predict a significant role of nominal shocks in the current account fluctuations for a number of OECD countries. Moreover, the estimated IRFs show that in almost all countries a positive monetary shock generates a temporary and statistically significant current account surplus. Both results support the empirical relevance of the basic monetary interterporal *Redux* model.

Degree of openness and current account fluctuations

In this Section, we focus on an additional prediction of the OR model which refers to relationship between country size and current account effects of monetary shocks. In particular, by focusing on the two final equations (2.15) and (2.17) of the basic framework, it can be seen that the larger (smaller) the home country – that is, the greater (the smaller) n – the less (the more) the positive impact of a home money increase on its current account (Obstfeld and Rogoff, 1995, page 643). As a result, we can expect that smaller and more open countries will show relatively greater effects of monetary shocks on the current account. Table 1 provides two often used indicators of openness to international trade: the

[31] In the U.S. case, Lane (1999) provides similar results and shows that the identified monetary shock accounts for about 50% of the variance of the current account. His specification, however, does not provide similar results for the other countries. In particular, we have estimated the same trivariate VAR system (i.e. relative output, the ratio of current account to output and relative prices) and imposed the same long-run restrictions for the G7 over our sample period. Results are much less clear-cut in that, while the IRFs are qualitatively similar to ours, they are not statistically significant. Moreover, with the only exception of the U.S., the size of the contribution of monetary shocks in current account fluctuations is less important (i.e. between 5-10 per cent

ratio of total exports to real GDP and the ratio of total trade volumes to real GDP. Figures refer to average values over the period 1980-1998 GDP for the 15 OECD countries. From the two columns, it is evident that the two ratios provide the same classification. In particular, as expected, relatively smaller countries seem to be more open, whereas low values of the indicators are associated with relatively larger countries.

In order to test formally for this prediction, we standardise the maximum effect of the estimated current account impulse response to the identified monetary policy shock according for all the countries. Figure 2 plots this criterion against the ratio of the overall trade to real GDP. Although the specific proxy for the trade openness and the imprecise estimates of the impulse responses might bias such an exercise, the upward fitted line clearly shows a positive correlation. This result supports the theoretical prediction from above, and might be interpreted as an additional positive result demonstrating the consistency of the VAR system and the theory.

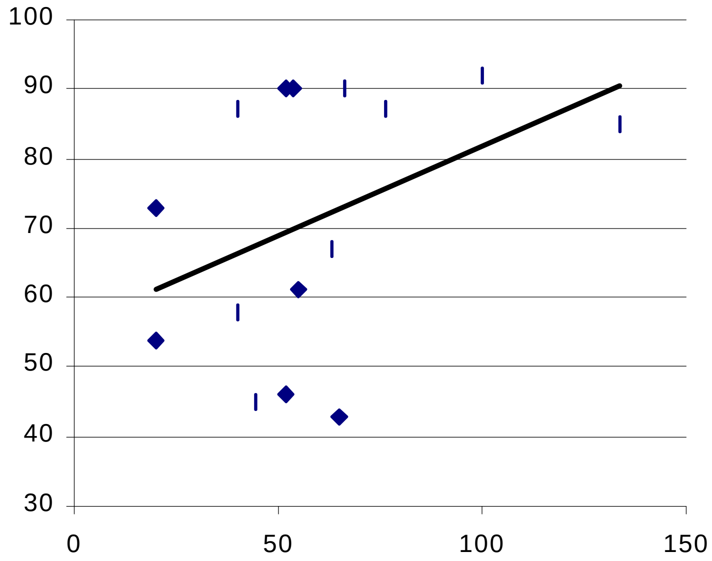
TABLE 1 Measures of Openness

	Overall exports as % GDP	***Overall exports plus imports as % GDP***
Austria	*37.1*	*74.0*
Belgium	*68.5*	*134.6*
Canada	*27.5*	*54.1*
Denmark	*33.9*	*65.2*
Finland	*28.8*	*56.2*
France	*22.0*	*43.4*
Germany	*26.5*	*50.8*
Italy	*19.9*	*39.9*
Netherlands	*11.4*	*100.6*
Portugal	*30.6*	*69.1*
Spain	*19.2*	*39.4*
Sweden	*32.2*	*62.5*
United Kingdom	*25.6*	*52.2*
United States	*9.0*	*19.3*

Sources: our estimates on OECD data.

over six years forecast horizon). These results might depend on the fact that a statistically incorrect long-run restriction is imposed on the current account to GDP ratio, which is stationary in our sample period.

CHART 1 Openness and current account effects to MS



Source: our estimates

7. Conclusions

The intertemporal approach to the current account has been of fundamental importance in explaining current account imbalances. However, most of the literature has focused on productivity and fiscal shocks as main determinants of its variations. In this paper, we emphasise the importance of nominal (or monetary) shocks in generating exchange rate and current account fluctuations, by first outlining the standard monetary version (and some closely related extensions) of the sticky price intertemporal model of Obstfeld and Rogoff (1995, 1996), and then empirically investigating the main predictions of this class of models.
We use two structural VAR models to estimate the responses to real and nominal shocks across 15 OECD countries over the period 1979Q1 to 1998Q4. Similarly to Blanchard and Quah (1989), we achieve identification of the model by imposing long-run restrictions and leaving the short run fluctuations of the variables to be determined by the data. The forecast error variance decomposition shows strong evidence in favour of a crucial role of monetary shocks in generating current account movements in most of the countries. Moreover, the estimated impulse response functions to the identified monetary shock are very much in line with the main predictions of the Obstfeld and Rogoff model, in which unexpected monetary policy shocks temporarily and significantly improve the current account. Our results provide strong support in favour of the basic *Redux* model, and weak empirical evidence of those extensions of the basic framework – introduction of high degree of PTM behaviour (Betts and Devereux, 2000) and poor elasticity of substitution between domestic and imported goods (Tille, 2001)- in which a current account neutrality or deficit is predicted.
In the last Section we have verified a further prediction of this class of models. In particular, we have shown that trade openness of the country, which is strictly dependent on the size of the 15 economies under investigation, is positively correlated with the estimated maximum effect of the current account to the identified monetary shock.

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APPENDIX 1 – Data Description and Unit Root Analysis

As in Lee and Chinn (1998), in the first bivariate system we use the CPI-deflated real exchange rate, the nominal current account in US dollars, and the national-currency denominated GDP in nominal terms data taken from the International Financial Statistics (IFS) published by the International Monetary Fund.[32] Additionally, for the second specification, we take seasonally adjusted GDP data denominated in national currency taken from the same source. In our estimations, we use the log of the real exchange rate (REER), the ratio of the current account to GDP (CA/Y), and the log ratio of real domestic GDP to foreign real GDP (Y/Y*).

The original current account series are in US dollars and non-seasonally adjusted. Following Lee and Chinn (1998), we obtain the national currency-denominated current account, by using the average bilateral exchange rate of each period. We then calculate the ratio CA/Y, and seasonally adjust them using dummy variables. For the G7 countries, the rest of the world (or foreign) real GDP for each country is proxied and constructed as the trade-weighted sum of US$-denominated real GDP of the remaining G7.

For most of the countries, the data are over 1979:1-1999:4 (1979:1-1998:4 in the three-variable system). The only exceptions are Belgium and Denmark, in which data are only available after 1985:1 and 1987:1, respectively.

Testing for Unit Roots

Tables 1, 2 and 3 tabulate the unit root test results for all of the series. Column 1 shows the ADF test statistics. Column 2 provides the optimum lag length, which has been selected on the basis of the significance of the coefficients of the lagged terms, and the white noise properties of the residuals[33]. In the third column, we indicate the deterministic components in the regression model used for testing. Alternative deterministic term combinations have been tested in order to reinforce the reported results. In the last column, we show the critical values at the 1% and 5% level of significance provided by MacKinnon (1991).

As far as the REER series is concerned, the results suggest that we cannot reject the null hypothesis of non-stationarity indicating that all the variables are not stationary in levels.

[32] It is worth pointing out that, according to the Fund's definition and construction of the index, a decrease in the real effective exchange rate reflects a depreciation.

The only exceptions are France and the Netherlands, where we can reject the null hypothesis of the presence of a unit root at 5% significance level.[34] On the basis of this, we implement the same test for the first differences of the series. Although we do not show the results, the relevant ADF test statistics all reject the non-stationary hypothesis at any level of significance. Consistent with a number of other relevant studies on the post Bretton Wood period, we then assume that all REER series are first-difference stationary.

TABLE 1 Augmented Dickey-Fuller Test for the Presence of Unit Roots in the REER series (1979:1–1998:4)

Country	*ADF test statistics*	*ADF(p) lag length*	*Deterministic component*	*Critical values at 1% and 5% significance level**
Austria	-1.58	ADF(1)	const no trend	-3.51 and –2.89
Belgium	-2.73	ADF(1)	const no trend	-3.51 and –2.89
Canada	-1.19	ADF(3)	const no trend	-3.51 and –2.89
Denmark	-3.16	ADF(1)	const and trend	-4.07 and –3.46
Finland	-2.07	ADF(1)	const and trend	-4.07 and –3.46
France	-3.01(*)	ADF(1)	const no trend	-3.51 and –2.89
Germany	-2.75	ADF(1)	const no trend	-3.51 and –2.89
Italy	-1.94	ADF(1)	const no trend	-3.51 and –2.89
Japan	-2.59	ADF(3)	const and trend	-4.08 and –3.46
Netherlands	-3.24(*)	ADF(1)	const no trend	-3.51 and –2.89
Portugal	-1.96	ADF(1)	const and trend	-4.07 and –3.46
Spain	-1.66	ADF(2)	const no trend	-3.51 and –2.89
Sweden	-1.94	ADF(1)	const no trend	-3.51 and –2.89
U.K.	-2.42	ADF(1)	const no trend	-3.51 and –2.89
U.S.	-1.52	ADF(1)	const no trend	-3.51 and –2.89

Notes: All variables are in logs. ADF(k) provides the optimum lag length k. The critical values are from MacKinnon. (*) and (**) indicate rejection of the null hypothesis of unit root at the 95% and 99% confidence level, respectively.

Table 2 shows the results we have found for CA/Y. For most of the countries, the stationarity of our current account series cannot be rejected at 0.05 significance level. The only exceptions are France, Germany and Sweden where the unit root hypothesis cannot be rejected at 10% level. Given the low power of the test, we check for alternative specifications of the ADF regressions, which allow us not to reject the stationarity of the above three series. On the basis of these results, we decide to implement an additional unit

[33] The presence of serially-correlated residuals has been tested by using the Lagrange multiplier (LM) test up to the third order as well as the Durbin-Watson test for first-order serial correlation.

[34] However, the test statistics are strongly dependent on the chosen specification. By testing for alternative number of lags and different deterministic components, we cannot reject the nonstationarity hypothesis at 5% level.

root test, that is, the KPSS test. Our results (not shown) are again supportive of the hypothesis of stationarity of the series.[35]

TABLE 2 Augmented Dickey-Fuller Test for the Presence of Unit Roots in the CA/Y series (1979:1–1998:4)

Country	*ADF test statistics*	*ADF(p) lag length*	*Deterministic component*	*Critical values at 1% and 5% significance level**
Austria	-3.85(**)	ADF(0)	no const no trend	-2.59 and –1.94
Belgium	-5.68(**)	ADF(0)	const and trend	-4.13 and –3.59
Canada	-2.96(**)	ADF(0)	no const no trend	-2.59 and –1.94
Denmark	-2.09(*)	ADF(3)	no const no trend	-2.61 and –1.94
Finland	-2.06(*)	ADF(0)	no const no trend	-2.59 and –1.94
France	-3.16	ADF(1)	const and trend	-4.08 and –3.46
Germany	-1.86	ADF(1)	no const no trend	-2.59 and –1.94
Italy	-2.48(*)	ADF(0)	no const no trend	-2.59 and –1.94
Japan	-2.35(*)	ADF(0)	no const no trend	-2.59 and –1.94
Netherlands	-2.02(*)	ADF(2)	no const no trend	-2.59 and –1.94
Portugal	-2.00(*)	ADF(1)	no const and trend	-2.59 and –1.94
Spain	-2.14(*)	ADF(2)	no const and trend	-2.59 and –1.94
Sweden	-1.83	ADF(1)	no const no trend	-2.59 and –1.94
U.K.	-1.95(*)	ADF(1)	no const no trend	-2.59 and –1.94
U.S.	-2.17(*)	ADF(6)	no const no trend	-2.59 and –1.94

Notes: All variables are in logs. ADF (k) provides the optimum lag length k. the critical values are from MacKinnon. (*) and (**) indicate rejection of the null hypothesis of unit root at the 95% and 99% confidence level, respectively.

Finally, Table 3 provides the results for the relative output variable Y/Y*. For all countries, we cannot reject the null hypothesis of non-stationarity at any level of significance. By implementing the same test for the first differences of the series (results not shown), the relevant ADF test statistics all reject the non-stationary hypothesis. Therefore, we conclude that the relative output Y/Y* is non-stationary in levels, and stationary in first differences.

[35] These results are consistent with the stationarity test on the same variable for the G-7 countries implemented by Lee and Chinn (1998). In their paper, they use the KPSS test and find that, for most countries, the stationarity of the current account series cannot be rejected at 5% level.

TABLE 3 Augmented Dickey-Fuller Test for the Presence of Unit Roots in the Y/Y* series (1979:1–1998:4)

Country	*ADF test statistics*	*ADF(p) lag length*	*Deterministic component*	*Critical values at 1% and 5% significance level*
Austria	-2.12	ADF(4)	const no trend	-3.51 and –2.89
Belgium	-2.06	ADF(1)	const no trend	-3.51 and –2.89
Canada	-0.79	ADF(1)	const and trend	-4.08 and –3.46
Denmark	-3.21(*)	ADF(1)	const no trend	-3.51 and –2.89
Finland	-1.38	ADF(1)	const no trend	-3.51 and –2.89
France	-2.34	ADF(1)	const no trend	-3.51 and –2.89
Germany	-1.96	ADF(1)	const and trend	-4.07 and –3.46
Italy	-2.09	ADF(1)	const and trend	-4.07 and –3.46
Japan	-1.53	ADF(1)	no const no trend	-2.59 and –1.94
Netherlands	-2.51	ADF(1)	const and trend	-4.07 and –3.46
Portugal	-1.85	ADF(1)	const no trend	-3.51 and –2.89
Spain	-2.67	ADF(1)	const no trend	-3.51 and –2.89
Sweden	-2.24	ADF(1)	const and trend	-4.07 and –3.46
U.K.	-2.14	ADF(1)	const no trend	-3.51 and –2.89
U.S.	-0.49	ADF(3)	no const no trend	-2.59 and –1.94

Notes: All variables are in logs. ADF(k) provides the optimum lag length k. The critical values are from MacKinnon. (*) and (**) indicate rejection of the null hypothesis of unit root at the 95% and 99% confidence level, respectively.

FIGURE 2 Impulse Responses for France. Bivariate VAR

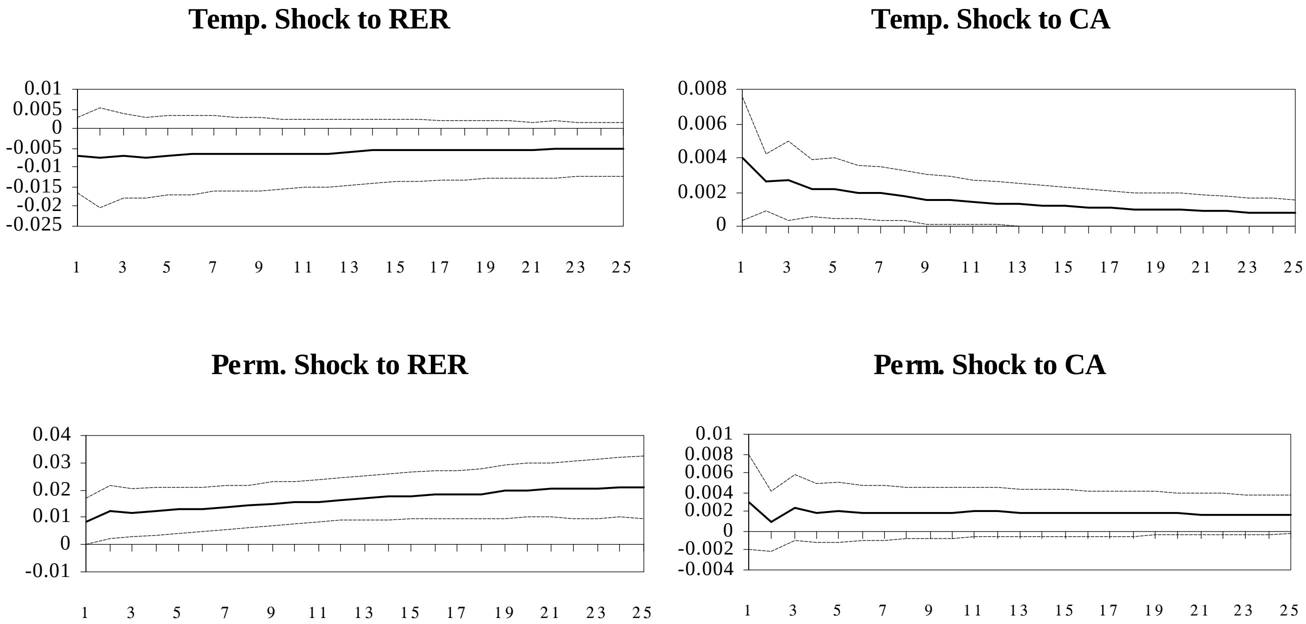


FIGURE 3 Impulse Responses for Spain. Bivariate VAR

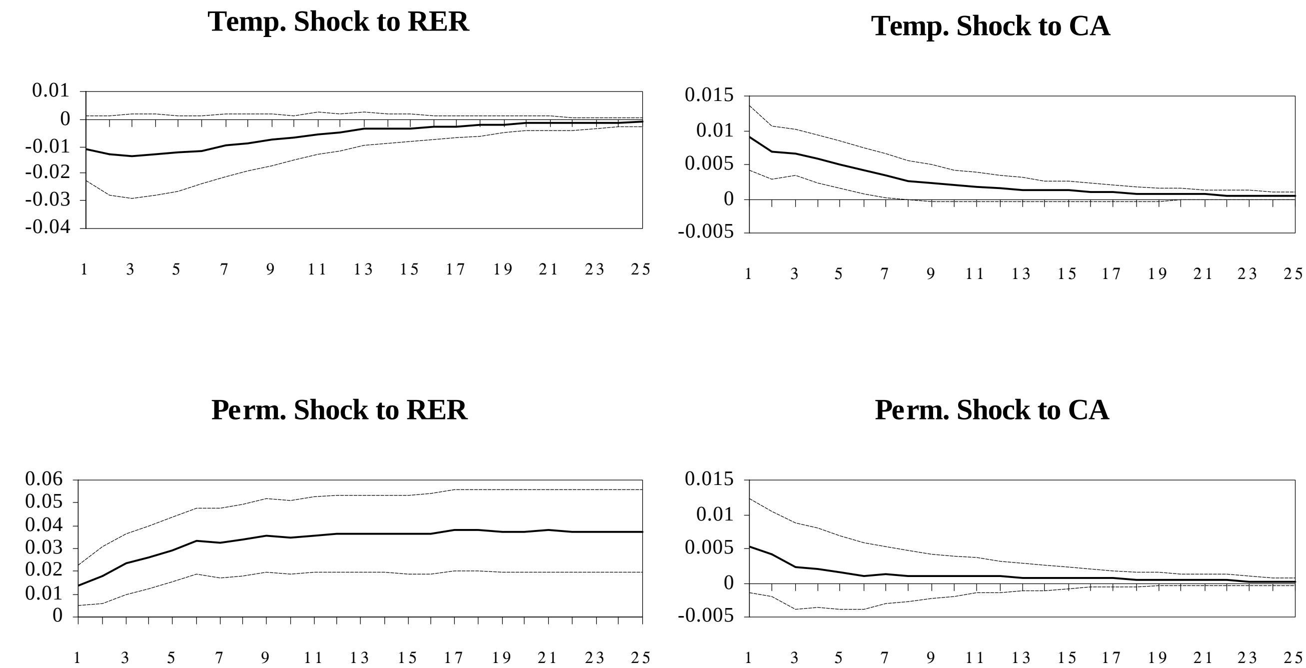


FIGURE 4 Impulse Responses for Sweden. Bivariate VAR

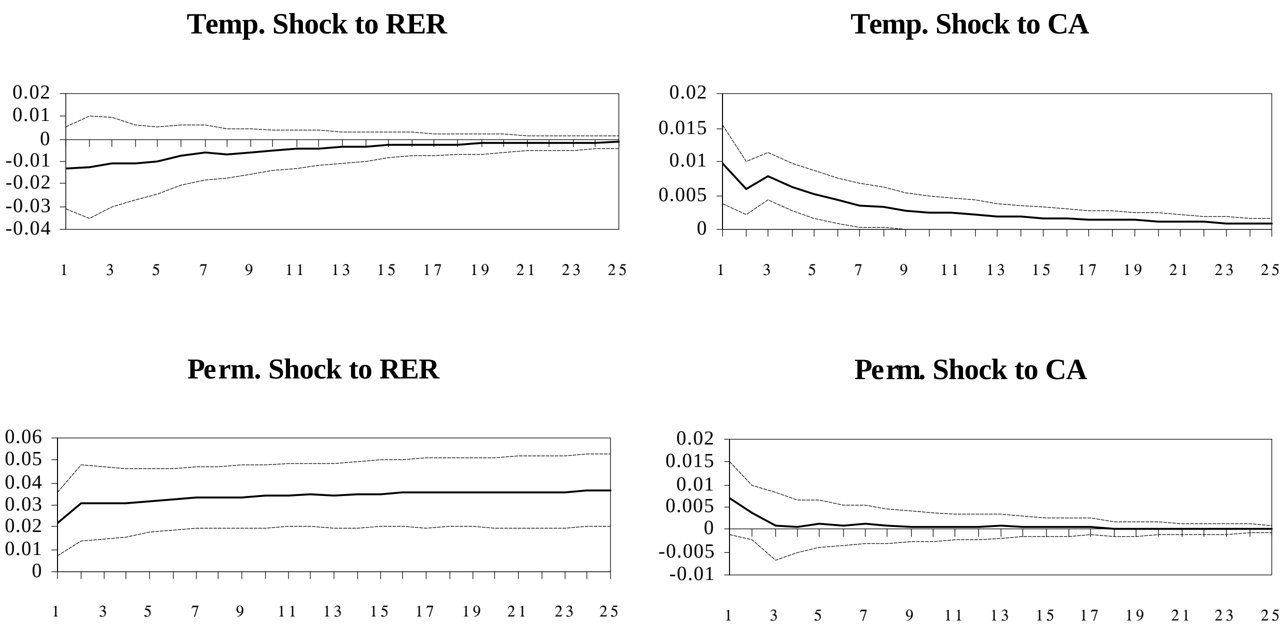


FIGURE 5 Impulse Responses for the U.S.. Bivariate VAR

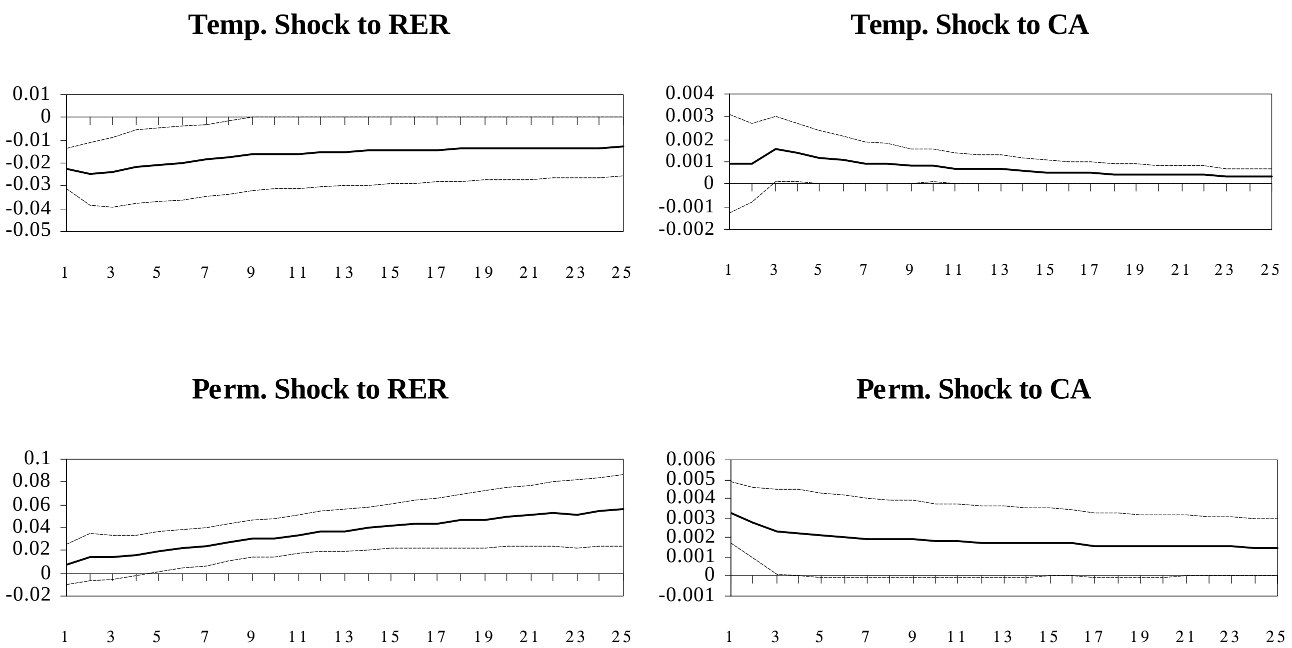


TABLE 4 Variance Decomposition of Current Account. Bivariate VAR.

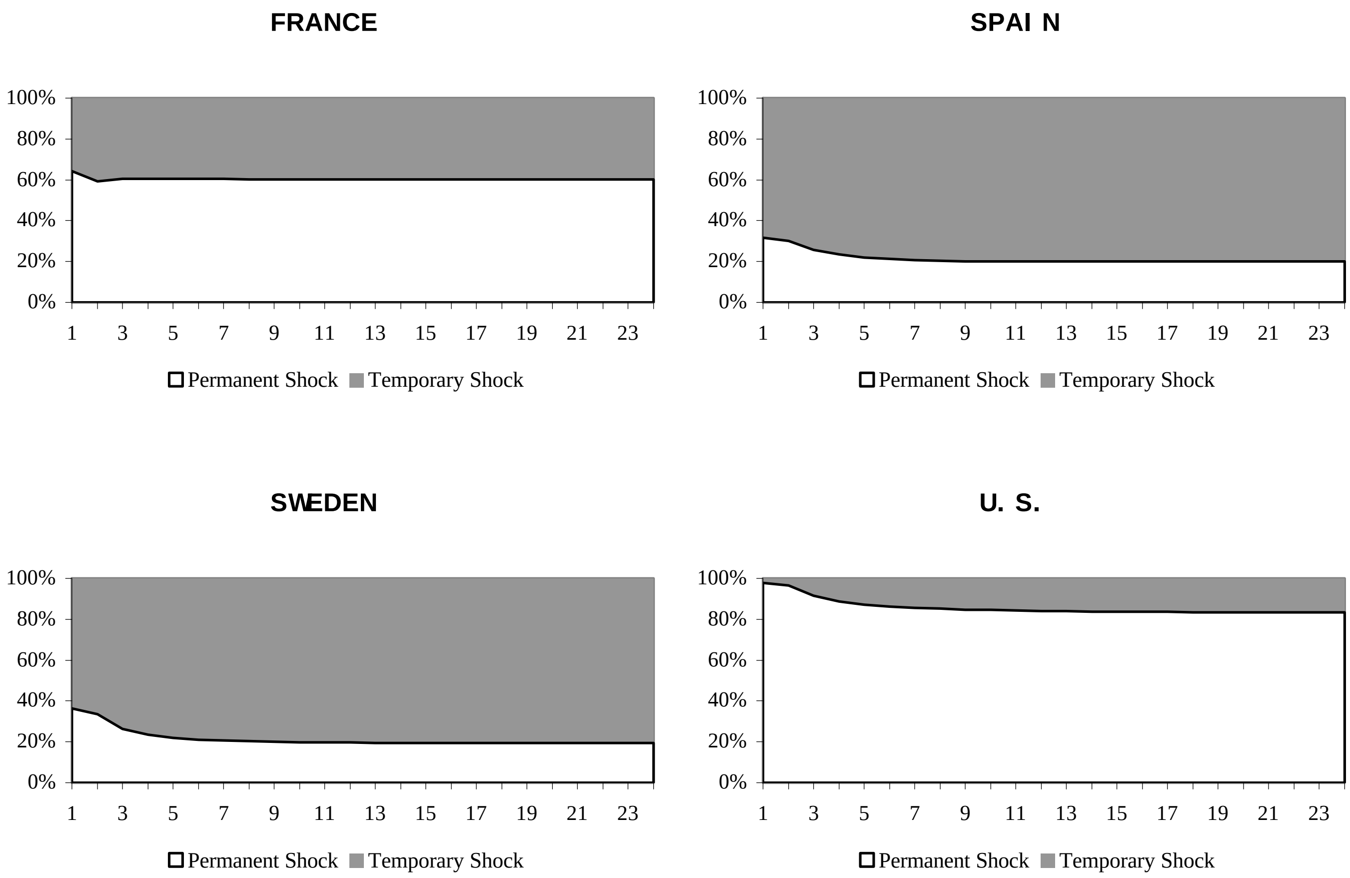
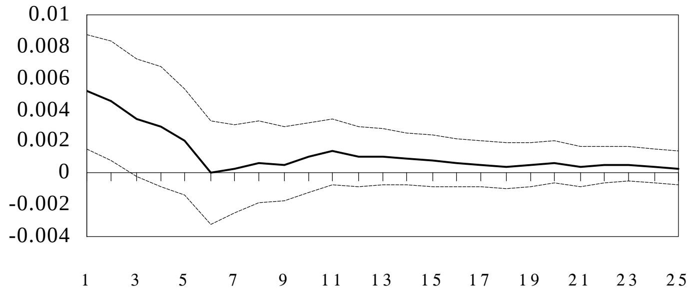
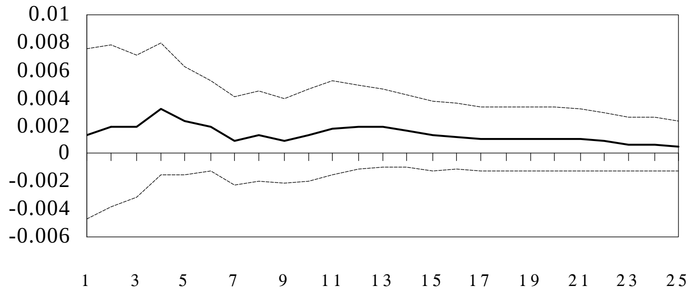


FIGURE 6 Response of CA to Structural Shocks for Canada. Trivariate VAR.

Monetary Shock



Demand Shock



Supply Shock

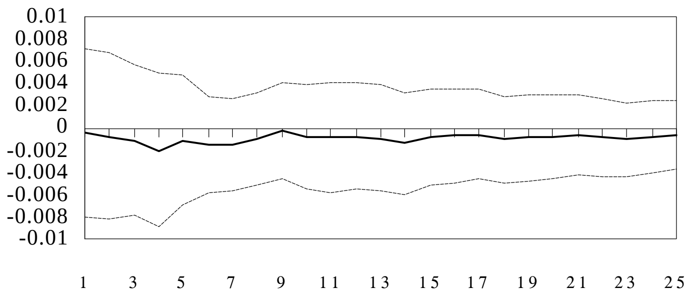
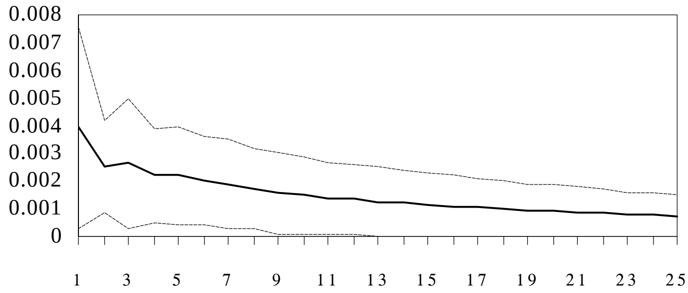
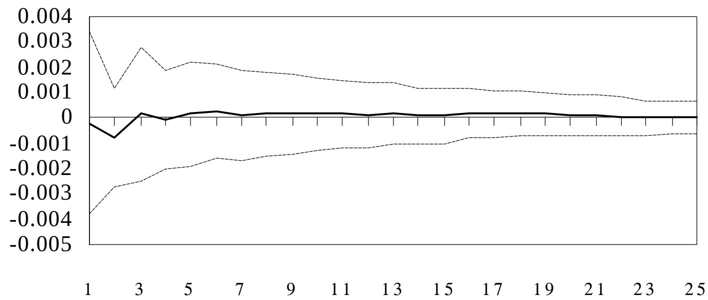


FIGURE 7 Response of CA to Structural Shocks for France. Trivariate VAR

Monetary Shock



Demand Shock



Supply Shock

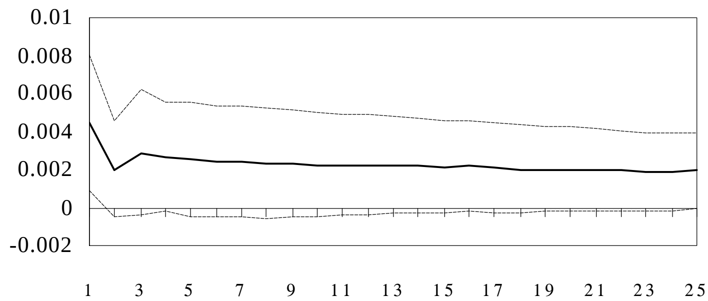
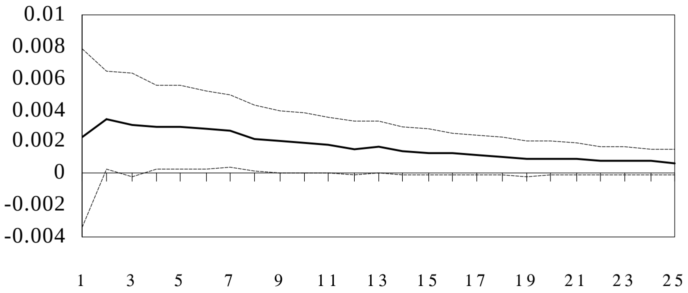
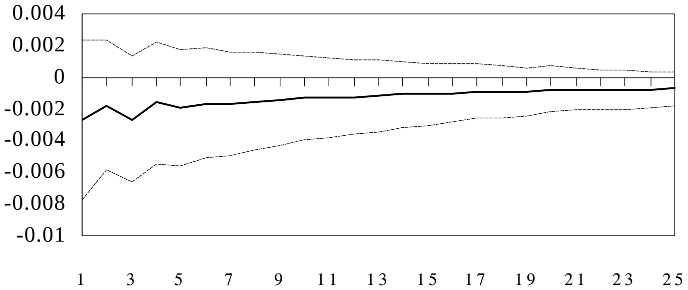


FIGURE 8 Response of CA to Structural Shocks for Germany. Trivariate VAR

Monetary Shock



Demand Shock



Supply Shock

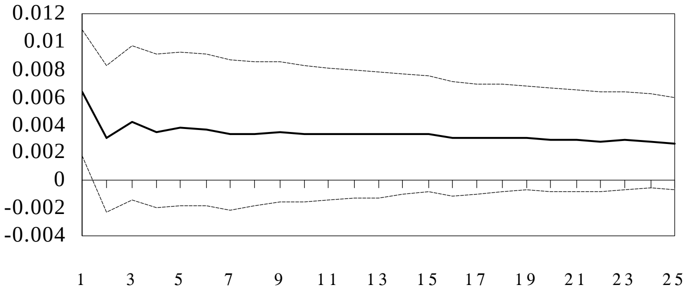
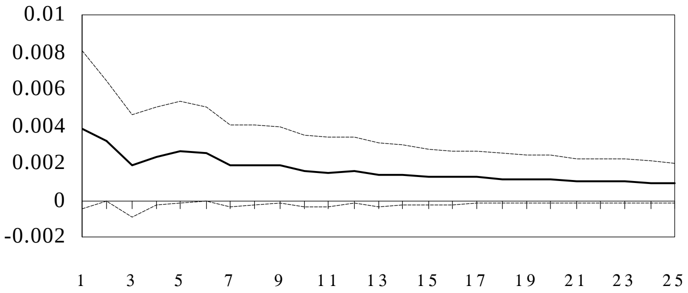
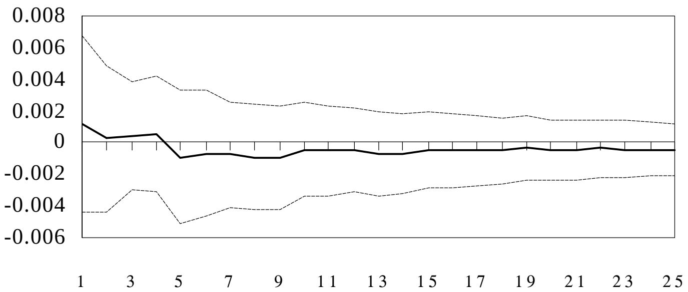


FIGURE 9 Response of CA to Structural Shocks for Italy. Trivariate VAR

Monetary Shock



Demand Shock



Supply Shock

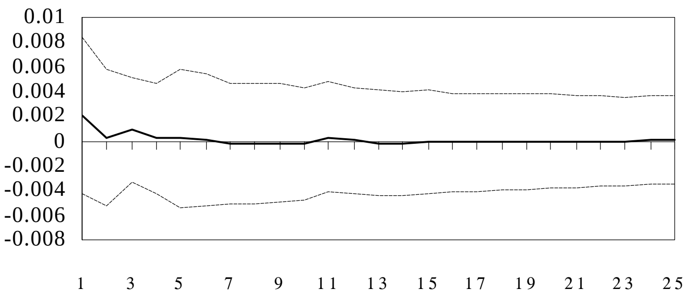
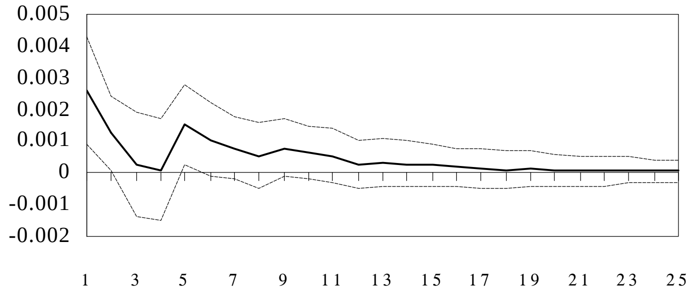
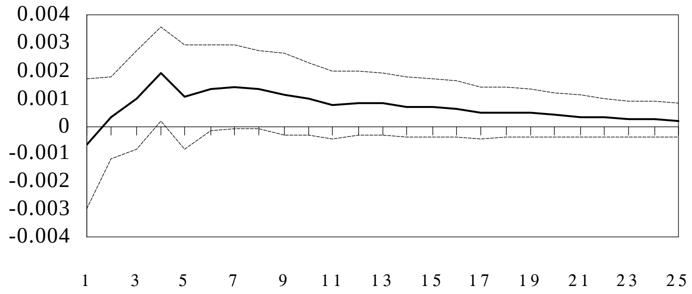


FIGURE 10 Response of CA to Structural Shocks for Japan. Trivariate VAR

Monetary Shock



Demand Shock



Supply Shock

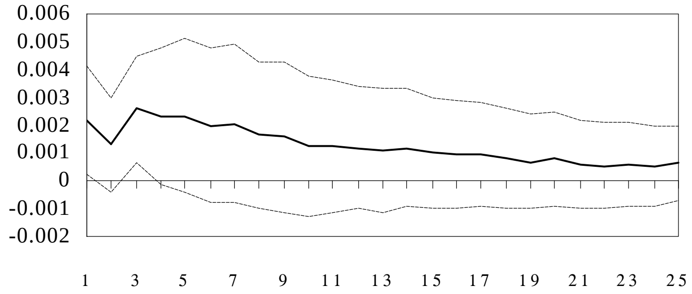
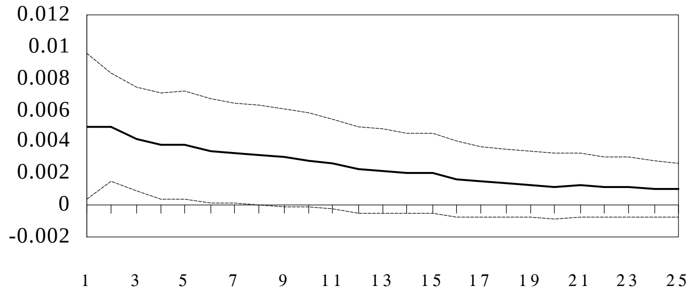
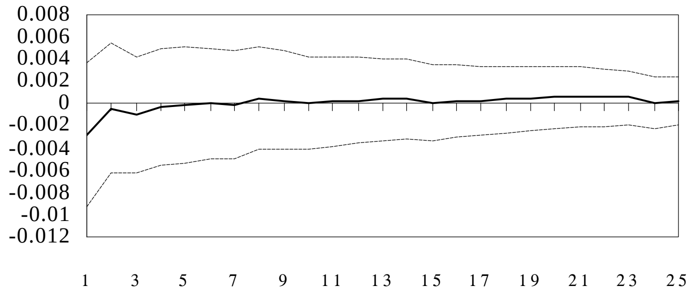


FIGURE 11 Response of CA to Structural Shocks for the U.K.. Trivariate VAR

Monetary Shock



Demand Shock



Supply Shock

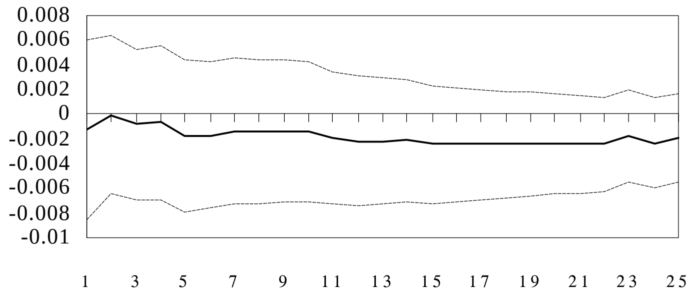
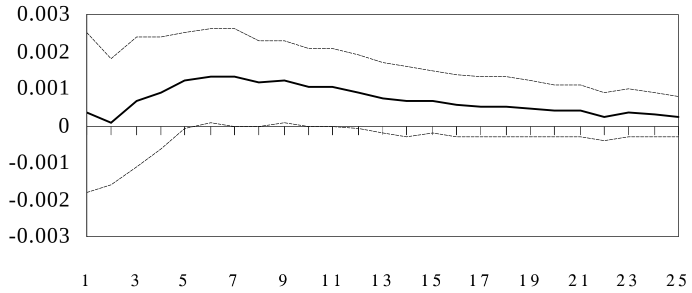
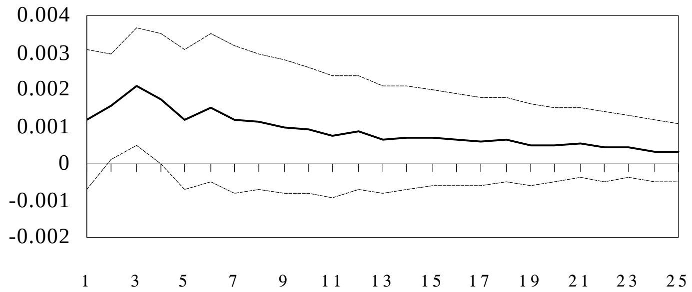


FIGURE 12 Response of CA to Structural Shocks for the U.S.. Trivariate VAR

Monetary Shock



Demand Shock



Supply Shock

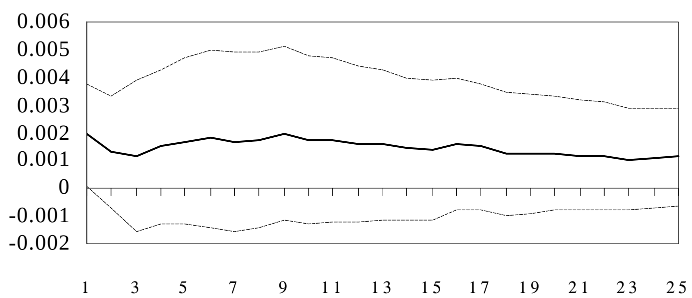
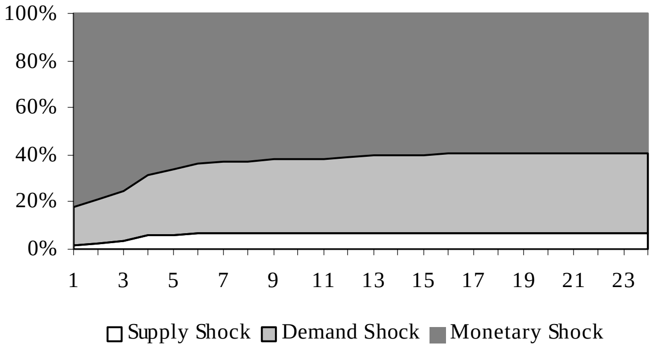
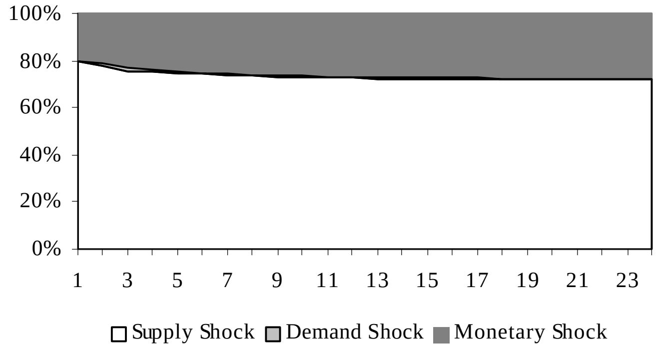


TABLE 4 Variance Decomposition of the Current Account. Trivariate VAR.

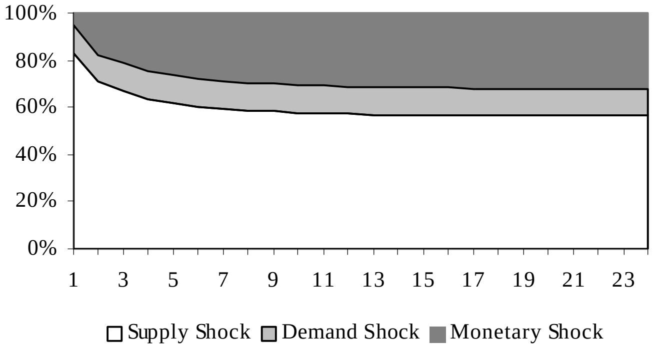
CANADA



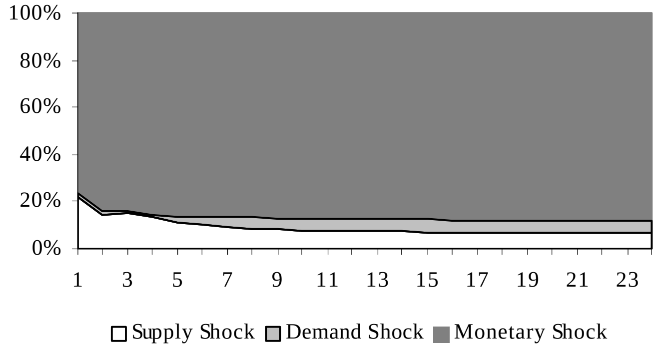
FRANCE



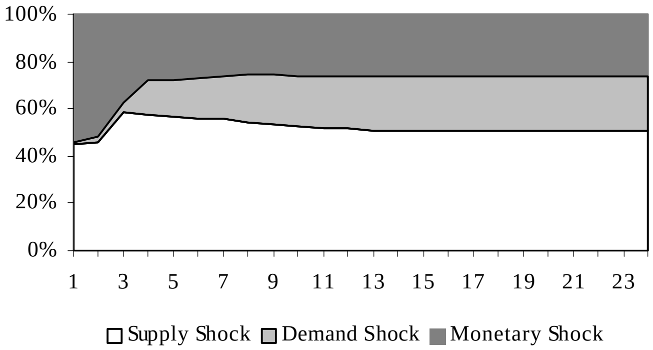
GERMANY



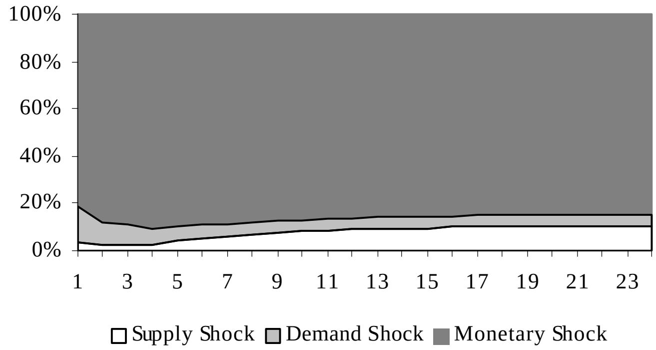
I TALY



JAPAN



U. K.



U. S.