FISCAL POLICY AND REAL EXCHANGE RATE. FISCAL IMPULSES OR INTERTEMPORAL APPROACH?*

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ABSTRACT

In this paper we compare the results of applying two differente approaches to the determination of real exchange rates. The first approach is based on relative fiscal impulses, as defined by the IMF, that account for demand shocks and was implicitly proposed by Obsfeld (1985). For the second approach, we follow a transformed version of Rogoff (1992) intertemporal neoclassical fixed-factors fiscal model for real exchange rate determination. We thus estimate a relationship between the peseta real exchange rate and productivity, government spending and the real price of oil

Keywords: fiscal policy, real exchange rate, cointegration
JEL: C32, F31, H61.

RESUMEN

En este trabajo se comparan los resultados de aplicar dos enfoques distintos para la determinación del tipo de cambio real de la peseta respecto a Alemania, Francia y el Reino Unido. El primer enfoque se basa en el modelo propuesto por Obsfeld (1985), y explicaría el tipo de cambio real en función de perturbaciones de demanda. Aquí se han utilizado los impulsos fiscales, siguiendo la definición del FMI, para recoger dichos shocks, tal y como se derivva implícitamente del planteamiento de Obsfeld. El segundo enfoque está basado en una versión transformada del modelo fiscal par ala determinación del tipo de cambio real de Rogoff (1992), de carácter intetemporal y que considera los factores productivos fijos en el corto-medio plazo. Este último modelo da lugar a una expresión que relaciona el tipo de cambio real con la productividad, el gato público y el precio real de la energía.

Palabras clave: política fiscal, tipo de cambio real, cointegración.
JEL: CFH

1 Introduction.

The last twenty years have witnessed some dramatic changes in the relationships linking the Spanish macroeconomic variables. Specifically, the fiscal policy has varied both its composition and its structure if compared to the policy applied in the seventies. Moreover, the share of the public sector in the economy has increased substantially, so that the rest of macroeconomic variables should also have been influenced[1].

At the same time, the peseta real exchange rate has sharply appreciated from the end of the last decade and the beginning of the nineties. This fact caused an important loss of competitiveness in the exports sector that finally led to a series of devaluations of the peseta from 1992 to 1995.

An additional fact has increased the general interest for the fiscal variables: the prospect of the European Monetary Union (EMU) and the loss of the monetary and exchange rate policies in order to affect the internal and the external balance of the economy. The Growth and Stability pact has set the conditions and the possible penalties that can suffer those countries that do not satisfy the two fiscal Maastricht criteria. Consequently, what it is required for the EMU member countries is to coordinate their fiscal policies and maintain the public finances within the limits to follow a sustainable path.

Camarero, Esteve and Tamarit (1996) studied some of the variables influencing the real effective exchange rate of the peseta relative to the European Union (EU). In this paper, they tried to determine the importance of the domestic policies implemented by monetary and fiscal authorities (represented by the interest rate differential) relative to the relative prices of trade (terms of trade). However, recent theoretical developments, such as Rogoff (1992), Obstfeld (1993), and Asea and Mendoza (1994), emphasize the role of fiscal policy and other real variables (productivity shocks, for example) in real exchange rate models, in contrast to the more traditional monetary approaches. In this context, Chinn (1997) compares the performance of two types of fiscal models to two conventional monetary models, for the determination of both nominal and real exchange rates.

This paper is organized as follows. The second section presents a Keynesian asset-based rational expectations model for real exchange rate determination proposed by Obstfeld (1985), which will be tested using the fiscal impulse variables to account for demand shocks as defined by the IMF. The third section presents an intertemporal fixed-factors model based on Rogoff

[1] See Martín (1997), chapter 5, for a detailed comparison of the European and Spanish fiscal policy and a development of this issues.

(1992). The forth section shows the main empirical results and briefly describes some of the data and techniques employed while the fifth section concludes.

2 Keynesian fiscal models and the fiscal impulses.

2.1 Theoretical model.

Chinn (1997) presents a Keynesian asset-based rational expectations model based on Obstfeld (1985). In this model, aggregate demand in both the domestic and the foreign country are the following:

$$y_t^d = \delta q_t - \sigma r_t + \gamma y_t^* + f_t \tag{1}$$

$$y_t^{*d} = -\delta^* q_t - \sigma^* r_t^* + \gamma^* y_t + f_t^* \tag{2}$$

where y is national output, q the real exchange rate (both in logarithms), r is the domestic real interest rate, f an aggregate demand shock (for example, cyclically adjusted fiscal deficit), and asterisks denote foreign variables.

The corresponding equations for aggregate supply are the following:

$$y_t^s = \theta(p_t - w_t) \tag{3}$$

$$y_t^{*s} = \theta^*(p_t^* - w_t^*) \tag{4}$$

where wages are predetermined at t and, consequently, expected output is set as potential.

The money demand equations are described as:

$$m_t - p_t = \varphi y_t - \lambda i_t \tag{5}$$

$$m_t^* - p_t^* = \varphi^* y_t^* - \lambda^* i_t^* \tag{6}$$

In the asset market, the fulfillment of uncovered interest parity reflects the existence of international linkages:

$$i_t = i_t^* + E_t s_{t+1} - s_t \tag{7}$$

Given also the fulfillment of the Fisherian definition of the *ex ante* real interest rate, the *ex ante* real interest rate differential equals the expected real exchange rate depreciation:

$$r_t - r_t^* = E_t q_{t+1} - q_t \tag{8}$$

Finally, the expression for nominal exchange rate is the following:

$$s_t = m_t + \lambda\mu - m_t^* - \lambda^*\mu^* + q_t \qquad for \quad \lambda = \lambda^* \tag{9}$$

where μ is the trend growth rate of nominal money.

Solving the model, the real exchange rate is given by the expression:

$$q_t = \frac{1}{\sigma\sigma^* + \delta\sigma^* + \delta^*\sigma} \sum_{j=0}^{\infty} \left(\frac{\sigma\sigma^*}{\sigma\sigma^* + \delta\sigma^* + \delta^*\sigma} \right)^j (\sigma E_t f_{t+j}^* - \sigma^* E_t f_{t+j}) \tag{10}$$

In this rational expectations solution, the real exchange rate depends on the expected future path of its driving variables, which are, in this case, the fiscal shocks. However, equation (10) is not directly estimable. Assuming that each shock follows a random walk, it can be represented as:

$$q_t = \frac{\sigma}{\delta^*\sigma + \delta\sigma^*} f_t^* - \frac{\sigma^*}{\delta^*\sigma + \delta\sigma^*} f_t \tag{11}$$

As (11) is a long run relationship, the real exchange rate is cointegrated with real aggregate demand shocks. A positive external demand shocks depreciates the real exchange rate, while a domestic shock tends to appreciate it. However, real interest rates do not appear in the equation above because they are endogenous variables and have been substituted.

2.2 The IMF fiscal impulse measure.

In Obstfeld (1985), the author referred to this proxy, although he did not use it in the empirical analysis, mainly due to the short span of data available at that time. In the empirical analysis of this paper, we use the fiscal impulse as it is calculated by the IMF and reported annually in the *World Economic Outlook*. This variable is measured as a percentage of GNP or GDP, where a positive sign denotes an expansionary effect.

According to Chand (1993), this measure aims at providing more accurate indications of budget impact that the ones provided by simply observing movements in the actual budget balance. Although simple and subject to criticisms, the IMF version of the fiscal impulse is useful in indicating the approximate directions of fiscal impact. This indicator is used to assess the annual contribution (expansionary, neutral, or contractionary) of budgets to aggregate demand.

The IMF fiscal impulse can be described as:

$$FI = (\Delta G - g_0 \Delta YP) - (\Delta T - t_0 \Delta Y) \tag{12}$$

where FI stands for fiscal impulse, G for government expenditure, g_0 is the base-year ratio of G relative to potential gross national product (GNP), YP, T is government revenue (taxes) and t_0 is the base-year ratio of government expenditure to actual GNP, Y.

Chand (1993) derives the FI from the *cyclically neutral budget model* and the subsequent distinction between the changes in government revenue and expenditure associated with cyclical fluctuations in the output of an economy and the changes that reflect policy decisions. He defines the *cyclical effect of the budget (CEB)*, that can be obtained by substracting from the actual budget deficit for any year, the budget deficit that is cyclically neutral for that year:

$$CEB = (G - T) - (g_0 YP - t_0 Y) \tag{13}$$

The last right hand term in equation (13) is the cyclically neutral balance. The fiscal impulse of equation (12) has been derived taking first differences in (13) and rearranging. This indicator is more robust than the CEB because it does not depend on the chosen base year. In fact, the IMF's FI is a good approximation of an alternative measure of fiscal impulse, that is called *Dutch fiscal impulse*, where the impulse is determined by the preceding year's budget balance as a base[2]:

$$FI = (\Delta G - nG_{-1}) - [\Delta T - (\Delta Y / Y_{-1}) T_{-1}] \tag{14}$$

where $n = \Delta YP / YP_{-1}$ is the rate of growth of potential output. If one divides by the previous year's GNP and rearranging, can be derived another expression for the fiscal impulse:

[2]The equivalences are shown in Chand (1977).

$$FI/Y_{-1} = (\Delta G/G_{-1} - n)g^* - (\Delta T/T_{-1} - \Delta Y/Y_{-1})t^* \qquad (15)$$

where $g^* \equiv G_{-1}/T_{-1}$ and $t^* \equiv T_{-1}/Y_{-1}$ are the shares of government expenditure and revenue in the previous year's GNP, respectively.

Both equations (12) and (15) are equivalent, and this indicator involves testing the actual change in the budget deficit against a normative neutral change, as given by the movement in the cyclically neutral budget. If the actual change is larger than the normative change, the fiscal impulse is viewed as *expansionary.* It could be the result of excessive growth in expenditure or not enough growth in revenue, or a combination of the two. Its main advantage is that it can be considered a rate of growth (the initial fiscal contribution to the growth in aggregate demand) and can be easily calculated. Although it has been also criticized by Blanchard (1990), among others, other measures derived analytically are much more complex and do not have the advantage of simplicity and are not summary measures. Chand (1993) provides, however, a model-based rationale for the fiscal impulse indicator, based on a simple IS model. His derivation supports the idea of the fiscal impulse measure as identifying the active effects of fiscal policy on aggregate demand and eliminating induced effects on the budget from the overall rates of growth of the budget variables.

3 An intertemporal model for real exchange rate determination.

Rogoff (1992) assesses the development of a strand of theoretical literature on dynamic micro-foundation-based models of the real exchange rate, which emphasize the effects of real factors, such as productivity, government spending, taxes and terms of trade[3]. However, Rogoff thinks that it has been paid very little attention to the formulation of testable hypothesis that could be implemented in the context of non-stationary real exchange rates[4].

There is an extensive literature of models for exchange rate determination that incorporate non-traded goods. This strand has as seminal papers the works of Balassa (1964) and Samuelson (1964), but also more recent ones such as Obstfeld (1993) and Asea and Mendoza (1994). All these models incorporate factor mobility between the traded and non-traded goods sectors

[3]See, for example, Obstfeld (1982) as a seminal paper and Frankel and Razin (1987) who discuss the effects of fiscal policy.

[4]Rogoff (1992) points out at the near random walk behaviour of the real exchange rate.

and explain the non-stationarity of the real exchange rate as a function of the differential productivity growth trends in the two sectors.

However, in the traditional models, the capital markets are open, so that agents can smooth their consumption of tradeables when transitory productivity shocks occur to this sector. Consequently, the intra-temporal price of traded and non-traded goods is smoothed as well. Rogoff (1992) considers the more realistic (at least, for some countries) case of relatively closed capital markets and factors that are not perfectly mobile across sectors. The predictions of the two models concerning productivity differentials are similar, but this is not the case of government expending shocks. In the model with perfectly mobile factors, these shocks would not affect relative prices, while in the Rogoff model, both aggregate demand and supply shocks can be important, due to the fact that public expenditure is biased towards non-traded goods. Therefore, its effects cannot be smoothed intertemporally.

Rogoff (1992) developed a model that would be appropriate to study countries which are in the process of liberalizing capital markets, such as Japan. This could also be the case of Spain. He finds that traded goods productivity shocks will not be helpful in forecasting the real exchange rate, while the past behaviour of government consumption spending does help.

He presents an intertemporal model of the real exchange rate and the current account that we will try to summarize paying attention to the aspects related to the non-stationarity of real exchanges rates. Rogoff (1992) also assumes that the country is small in the sense that it cannot affect the world's interest rate.

The economy produces two types of goods, traded *("T")* and non-traded *("N")*. The production functions, that represent the supply side of the economy, are Cobb-Douglas:

$$Y_{Tt} = A_{Tt} L_{Tt}^{\theta T} K_{Tt}^{1-\theta T} \tag{16}$$

$$Y_{Nt} = A_{Nt} L_{Nt}^{\theta N} K_{Nt}^{1-\theta N} \tag{17}$$

where Y_T and Y_N are output in the traded and nontraded goods sectors, respectively. L, K and A are labour, capital and stochastic productivity shocks. It is assumed that capital and labour are fixed within each sector, so that there is no inter-sectorial mobility.

The demand side of the economy is represented by the time-separable utility function of the representative agent:

$$V_t = E_t \sum_{s=0}^{\infty} \beta^{s-t} \left[\frac{(C_{Ns}^{\alpha} C_{Ts}^{1-\alpha})^{1-\gamma}}{1-\gamma} \right] \tag{18}$$

where β is the subjective rate of discount, C_{Tt} and C_{Nt} are, respectively, the period-t consumption of tradables and non-tradables and γ the inverse of the elasticity of intertemporal substitution.

It should be noted that both government and private agents have free access to world capital markets, where they can borrow at gross interest rate R (measured in terms of tradables, so that agents can only save and borrow through trade in tradables). Ricardian equivalence holds as well as the intertemporal budget constraint. Since, by definition, non-traded goods cannot be exchanged intertemporally, domestic consumption of non-traded goods must equal output, while traded goods consumption can be smoothed intertemporally.

The relative price of non-traded goods in terms of traded goods, $\tilde{P}_t$, each period must depend on the relative consumption of the two goods:

$$\tilde{P}_t = \frac{\alpha C_{Tt}}{(1-\alpha) C_{Nt}} \tag{19}$$

In this model, it can be used interchangeably the terms "real exchange rate" and "relative price of nontraded goods". Additionally, the relative price of imports and exports (the terms of trade) are assumed constant, so that changes in the relative prices of non-traded goods are the only source of fluctuations in the real exchange rate.

The first order conditions imply that agents smooth expected marginal utility over time. Assuming homoskedasticity of the underlying productivity shocks, it can be approximated as:

$$E_t(c_{Tt+1} - c_{Tt}) \cong \frac{\alpha(1-\gamma)}{\gamma + \alpha(1-\gamma)} E_t(c_{Nt+1} - c_{Nt}) \tag{20}$$

where $\beta R = 1$. Taking logarithms in (19) one obtains $\tilde{p}_t = c_{Tt} - c_{Nt} + \log(\alpha/(1-\alpha))$ and combining it with (20) Rogoff shows that:

$$\tilde{p}_{t+1} - \tilde{p}_t = (c_{Tt+1} - c_{Nt+1}) - (c_{Tt} - c_{Nt}) \tag{21}$$

Equations (20) and (21) imply that barring shocks to the supply of non-traded goods available for private consumption, the log real exchange rate would follow a random walk independently of the serial correlation properties of the shocks to traded goods productivity.

The assumption that all factors are sector-specific can only be maintained in the short and maybe the medium run. However, Rogoff emphasizes the fact that even if shocks to traded goods are highly transitory they will induce long-lasting movements in the real exchange rate.

Concerning productivity shocks, and in order to obtain a testable expression, it is assumed that the shocks to both traded and non-traded goods productivity are lognormally distributed with homoskedastic disturbance terms:

$$a_{Nt+1} = \varphi a_{Nt} + \varepsilon_{Nt} \tag{22}$$

$$a_{Tt+1} = \rho a_{Tt} + \varepsilon_{Tt} \tag{23}$$

where $0 \leq \varphi, \rho \leq 1$, and the ε_t are independent disturbance terms. If one assumes that government spending, that falls solely on non-traded, follows a random walk, it is possible to obtain an expression for the first difference of the relative price of nontraded goods:

$$\tilde{p}_{t+1} - \tilde{p}_t = (a_{Tt+1} - a_{Tt}) - \zeta_N(a_{Nt+1} - a_{Nt}) + (\zeta_N - 1)(g_{Nt+1} - g_{Nt}) \tag{24}$$

where ζ_N is the ratio of non-traded goods output to private non-traded goods consumption. The larger ζ_N is, the greater the share of nontraded production accounted by the government sector. Changes in government spending would, thus, affect the price of non-traded goods and, hence, the real exchange rate, in proportion to that parameter.

Equation (24) can be rewritten in order to obtain a long-run cointegrating relationship, substituting backwards:

$$\tilde{p}_t = a_{Tt+1} - \zeta_N a_{Nt+1} + (\zeta_N - 1)g_{Nt+1} + \tilde{p}_0 \tag{25}$$

where p_0 includes the initial conditions. This equation relates the real exchange rate in one country to productivity in the traded and nontraded goods sectors and government spending. In order to extend the model to the relative price of two currencies, one can assume an identical foreign country. Substracting one from the other, it yields:

$$\begin{aligned} \tilde{p}_{t+1} - \tilde{p}^*_{t+1} &\equiv (p_{Nt+1} - p_{Tt+1}) - (s_t + p^*_{Nt+1} - s_t - p^*_{Tt+1}) \\ &\equiv \hat{a}_{Tt+1} - \zeta_N \hat{a}_{Nt+1} + (\zeta_N - 1)\hat{g}_{Nt+1} + \hat{p}_0 \end{aligned} \tag{26}$$

where the sign "^" denotes the difference between the domestic and foreign variables. If one assumes that PPP holds for the tradables goods, it is possible to write:

$$(s_{t+1} + p^*_{Nt+1} - p_{Nt+1}) = \hat{a}_{Tt+1} - \zeta_N \hat{a}_{Nt+1} + (\zeta_N - 1) g_{Nt+1} + \hat{p}_0 \qquad (27)$$

The real exchange rate is defined as:

$$\begin{aligned} q_t &\equiv (s_t + p^*_t - p_t) \\ &= \Gamma(s_t + p^*_{Nt} - p_{Nt}) \end{aligned} \qquad (28)$$

where Γ is the geometric weight of non-tradables in the aggregate price index. Substituting equation (27) into equation (28) results in:

$$q_{t+1} = -\Gamma[\hat{a}_{Tt+1} - \zeta_N \hat{a}_{Nt+1} + (\zeta_N - 1)\hat{g}_{Nt+1} + \hat{p}_0] \qquad (29)$$

Rogoff (1992) and Chinn (1997) use this expression to test the model in which the real exchange rate depends on the relative productivity in the traded and nontraded goods sectors, as well as the relative government spending (as a percentage of GDP). Rogoff (1992) also recommends to include the real price of oil to account for possible shifts in the terms of trade[5].

4 Empirical results.

4.1 Model based on fiscal impulses.

4.1.1 The Spanish fiscal policy: did Spain follow the same fiscal policy than the other European countries?

The approach proposed in the Keynesian model by Chinn (1997) relates the real exchange rate to domestic and foreign demand shocks. The variable that has been chosen to represent those shocks is the *fiscal impulse* as calculated by the IMF in its World Economic Outlook[6]. This technique involves a

[5] It has been assumed that the terms of trade are constant, so that the inclusion of the real price of oil permits to consider a possible source of exogenous shocks in the model.

[6] See Appendix A for further details.

distinction, with respect to government receipts and expenditures, between changes considered to be associated with cyclical fluctuations in an economy and other changes, which may be viewed as imparting expansionary or contractionary impulses to the economy. Consequently, this measure eliminates the more or less automatic responsiveness of government transactions to cyclical developments.

Previous to the application of the econometric techniques to the time series selected, Figure 1 shows the time path of the fiscal impulses in Spain, France, Germany and the United Kingdom. In order to make easier the visual inspection of the data, the variables are presented in pairs.

The fiscal impulses of Spain and France are presented in the upper left panel of Figure 1. In general, the two variables evolve in a similar way, although the Spanish fiscal impulses are larger in intensity. This means that both the expansions and the contractions are more important in volume in the Spanish case (see for example the observations of 1982 and 1985). There is one exception to this general behaviour of the public finances in the two countries: the period 1992-95, during which while the French fiscal authorities were expanding the Spanish ones contracted and conversely. Finally, in 1996 and 1997, both countries contracted, following the dictates of the Maastricht convergence criteria.

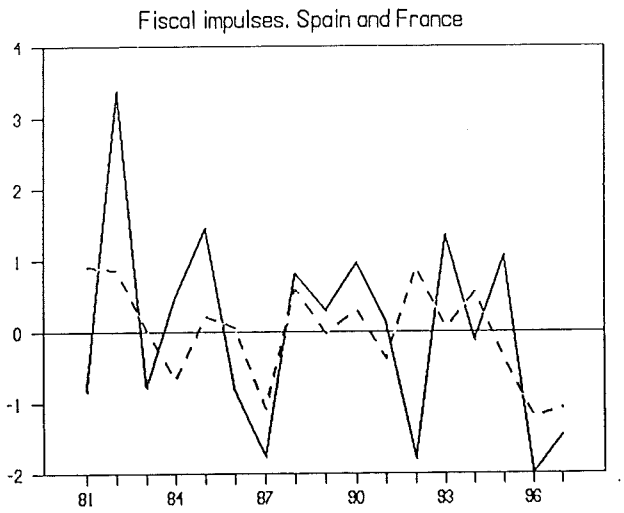
The upper right panel shows the fiscal impulses in Germany and Spain[7]. As in the previous case, the intensity of the fiscal impulses is, in general, larger in Spain. However, in contrast to the French fiscal policy, the German one seems to follow the opposite direction than the Spanish along the sample period. Only exceptionally (and again at the end of the sample, following the Maastricht criteria) the expansions and contractions coincide. A similar comment can be made about the evolution of the British fiscal impulses, as it is depicted in the lower left panel of Figure 1.

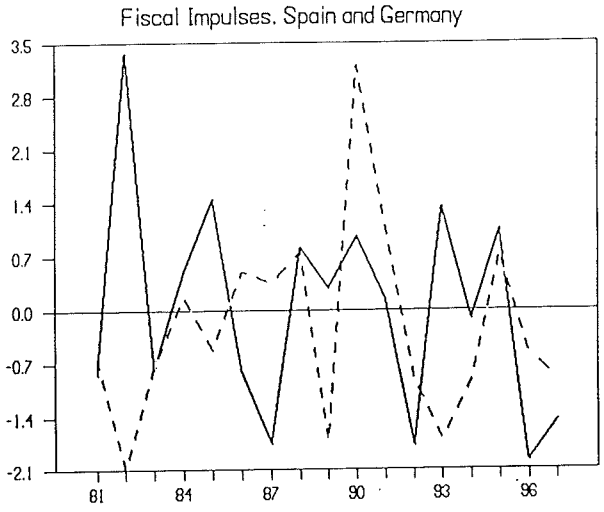
As a conclusion, the three graphs show that the evolution of the public finances in Spain have differed from those of the other three larger European countries analyzed. What can be even more surprising is the fact that, although Spain has followed the German setting of the monetary policy in the EMS discipline, the public sector has pursued its fiscal policy more or less independently. Only at the very end of the sample, once the Excess Deficit Procedure and the Maastricht convergence criteria have been working, the public accounts of the four countries have approached[8]. This fact makes us stress the emphasis given to the co-ordination of fiscal polices in a monetary

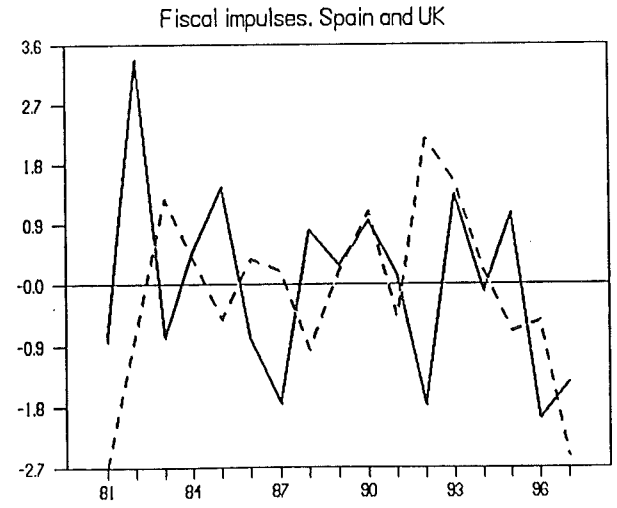
[7]It should be noted that the scale in the three graphs is not identical.

[8]Surprisingly, even the British economy, that has not participated in the process towards EMU has experienced negative fiscal impulses since the middle of the nineties.

FIGURE 1
FISCAL IMPULSES
SPAIN, FRANCE, GERMANY AND THE UK
1981-97







union with centralized monetary and exchange rate policy, where at least it should be guaranteed that each member state's fiscal policy is sustainable in the long-run.

4.1.2 Order of integration of the variables.

Previous to the analysis of the unit root tests, it should be given a note of caution. The fiscal impulse variables are only available for a very short span of data and on an annual basis. Chinn (1997) used a MA filter in order to generate quarterly data. According to Campbell and Perron (1991) this transformation would not add any substantial amount of information to the data and, probably, would simply introduce another filter that may pose some artificial problems. Consequently, we will proceed just using annual data and being conscious of the limited amount of information available.

Table 1 shows the unit root tests of the fiscal impulses and the bilateral real exchange rates. The Phillips-Perron tests allow us to easily reject the existence of two unit roots at the 1% level for all the series analyzed. However, two groups of variables can be distinguished for the null hypothesis of one unit root. On the one hand, for the fiscal impulses, it can be rejected again the existence of a unit root. On the other, none of the real exchange rates are stationary. The fourth and fifth columns of table 1 contain the KPSS unit root tests for the model with a time-trend (η_τ) and without it (η_μ). It should be emphasized that the null hypothesis of this test is the stationarity of the variable concerned. Thus, the rejection of the test is associated with a unit root. The results are very similar to those obtained using the Philips-Perron tests: the real exchange rates contain a unit root whereas the fiscal impulses are stationary, both in levels and in differences.

In order to complete the above results, we also run the Johansen unit root tests in the context of the cointegration relationships. Table 2 presents the stationarity tests for the three fiscal impulse models between Spain and Germany, France and the United Kingdom. According to this test, it is possible to reject the stationarity of the bilateral real exchange rates for the three cases, while it cannot be rejected the stationarity of the fiscal impulse in Spain. However, the fiscal impulses in Germany and the United Kingdom contain a unit root in the models with one cointegrating vector as well as in the model with two vectors for the cases of France and the UK. Consequently, the results are not as clear as in the Phillips-Perron tests but, in any case, one of the cointegrating vectors of each model, if there exist, should be the Spanish fiscal impulse, due to its stationarity.

C Tables.

TABLE 1
UNIT ROOT TESTS
FISCAL IMPULSE MODEL
1981-1997

Variables	Phillips-Perron			KPSS	
	$Z(t_{\tilde{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	η_μ	η_τ
	$l=1$			$l=1$	
Δfes_t	-9.09***	-9.23***	-8.84***	0.09	0.06
$\Delta fger_t$	-6.66***	-6.38***	-6.36***	0.06	0.05
$\Delta ffra_t$	-7.16***	-7.14***	-6.75***	0.10	0.09
Δfuk_t	-5.23***	-4.53***	-4.58***	0.31	0.07
$\Delta rerdm_t$	-3.80**	-3.65***	-3.63***	0.08	0.08
$\Delta rerff_t$	-4.08***	-3.91***	-3.88***	0.10	0.08
$\Delta rerps_t$	-7.09***	-5.46***	-5.44***	0.28	0.08
fes_t	-6.04***	-5.06***	-5.04***	0.21	0.06
$fger_t$	-3.89**	-3.91***	-3.80***	0.26	0.10
$ffra_t$	-3.63**	-3.24**	-3.25***	0.11	0.11
fuk_t	-2.83	-3.23**	-3.23***	0.13	0.13*
$rerdm_t$	-1.97	-1.87	-0.01	0.22	0.14**
$rerff_t$	-1.85	-1.78	-1.59	0.29	0.15**
$rerps_t$	-1.41	-2.00	-0.05	0.45**	0.20**

The critical values for the Phillips-Perron and KPSS tests are the following:

Signif. Level	$Z(t_{\tilde{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	η_μ	η_τ
1%	-4.38	-3.75	-2.66	0.73	0.21
5%	-3.60	-3.00	-1.95	0.46	0.14
10%	-3.24	-2.63	-1.60	0.34	0.11

in table 1, (*), (**) and (***) denote, respectively, rejection of the null hypothesis of a unit root at 10, 5 and 1% significance levels for the Phillips-Perron tests whereas they denote rejection of the null of stationarity for the KPSS test. The critical values for $Z(t_{\tilde{\alpha}})$, $Z(t_{\alpha^*})$ and $Z(t_{\hat{\alpha}})$ are taken from Fuller (1976), table 8.5.2 and the critical values for η_μ and η_τ appear in Kwiatkovski, Phillips, Schmith and Shin (1992), table 1.

TABLE 2
JOHANSEN TEST FOR STATIONARITY
FISCAL IMPULSE MODEL
1981-1997

Spain and Germany				
Coint. relations	$\chi^2(5)$	$rerdm_t$	$imes_t$	$imger_t$
1	5.99	16.03*	3.57	10.42*
2	3.84	8.71*	0.49	3.35
Spain and France				
Coint. relations	$\chi^2(5)$	rerfr_t	$imes_t$	$imfr_t$
1	5.99	6.98*	1.08	5.64
2	3.84	6.68*	1.08	5.27*
Spain and the UK				
Coint. relations	$\chi^2(5)$	$rerps_t$	$imes_t$	$imuk_t$
1	5.99	15.06*	1.09	13.32*
2	3.84	7.78*	0.72	5.69*

NOTE: The null hypothesis is stationarity. An asterisk denotes rejection of the null.

4.1.3 Cointegration results.

The previous section showed how the fiscal impulses turned out to be, in general, stationary. However, it would be still possible proceed and estimate the cointegrating relationships provided two out of the three variables involved were non-stationary. But previous to this analysis, the study of the time series properties of the variables did not recommend such possibility, due to the weak exogeneity[9] of the real exchange rates in this model. As it can be seen in table 3, except for the case of the Pound Sterling and two vectors, the exchange rates are not determined by the fiscal impulses. However, this analysis will be completed by means of Granger non-causality tests. This result is not striking, if we take into account that the unit root tests pointed to the stationarity of the IMF variables[10].

Chinn (1997) finds also, using Augmented Dickey Fuller tests, that the fiscal impulses of Germany and the United Kingdom are stationary, so that he is not able to estimate the Obstfeld model for the corresponding currencies.

4.1.4 Impulse-response functions and causal relations in the Obsfeld model.

Due to the stationarity of the fiscal impulse variables and the weak exogeneity of the real exchange rates, it is not possible to apply the cointegration techniques to the model proposed by Obstfeld (1985). Consequently, it won't be directly comparable to the intertemporal model. However, the use of a VAR model will allow us to determine the effects of innovations on the rest of the variables, specially the real exchange rate[11]. Three instruments will be applied to the variables of interest[12]: the impulse-response functions, the variance decompositions and the Granger test for non-causality[13].

The impulse-response functions. The impulse-response functions are plotted in figures 2, 3 and 4 for the German, French and British models,

[9] A definition of weak exogeneity in the context of the Johansen model is presented in appendix 2.

[10] Moreover, it should be taken into account that the definition of the fiscal impulses makes them rates of change, so that it is not surprising the absence of a unit root.

[11] In order to estimate a VAR model without cointegration restrictions, all the variables have to be stationary. Consequently, the real exchange rate is reported in differences, due to the fact that this variable was only stationary after differenciation.

[12] Due to the fact that the ordering of the variables may influence the results, the analysis has been applied following an order that starts from the most exogenous variable to the most endogenous one.

[13] See appendix B for details about the methodological aspects of this type of analysis.

TABLE 3
TESTS FOR WEAK EXOGENEITY
FISCAL IMPULSE MODEL
1981-1997

Spain and Germany				
Coint. relations	$\chi^2(5)$	$rerdm_t$	$imes_t$	$imger_t$
1	3.84	0.16	7.47*	0.03
2	5.99	1.45	16.95*	8.69*
Spain and France				
Coint. relations	$\chi^2(5)$	$rerff_t$	$imes_t$	$imfr_t$
1	3.84	0.31	0.45	0.23
2	5.99	2.46	10.18*	8.92*
Spain and the United Kingdom				
Coint. relations	$\chi^2(5)$	$rerps_t$	$imes_t$	$imuk_t$
1	3.84	2.55	7.67*	3.49
2	5.99	8.64*	16.39*	7.25*

NOTE: an asterisk denotes rejection of the null hypothesis of weak exogeneity.

FIGURE 2
IMPULSE-RESPONSE FUNCTIONS
FISCAL IMPULSE MODEL
SPAIN/GERMANY
1981-97

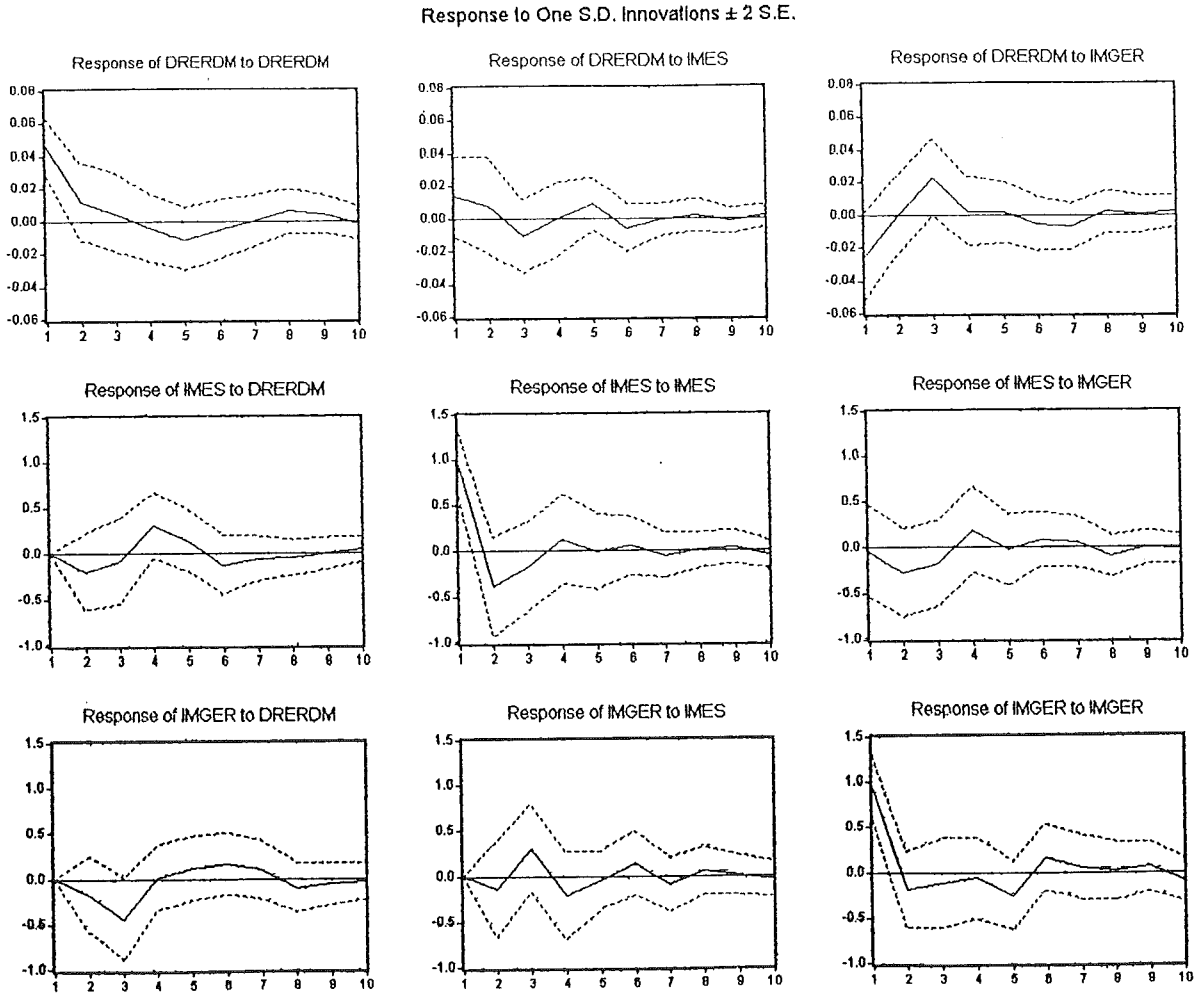


FIGURE 3
IMPULSE-RESPONSE FUNCTIONS
FISCAL IMPULSE MODEL
SPAIN/FRANCE
1981-97

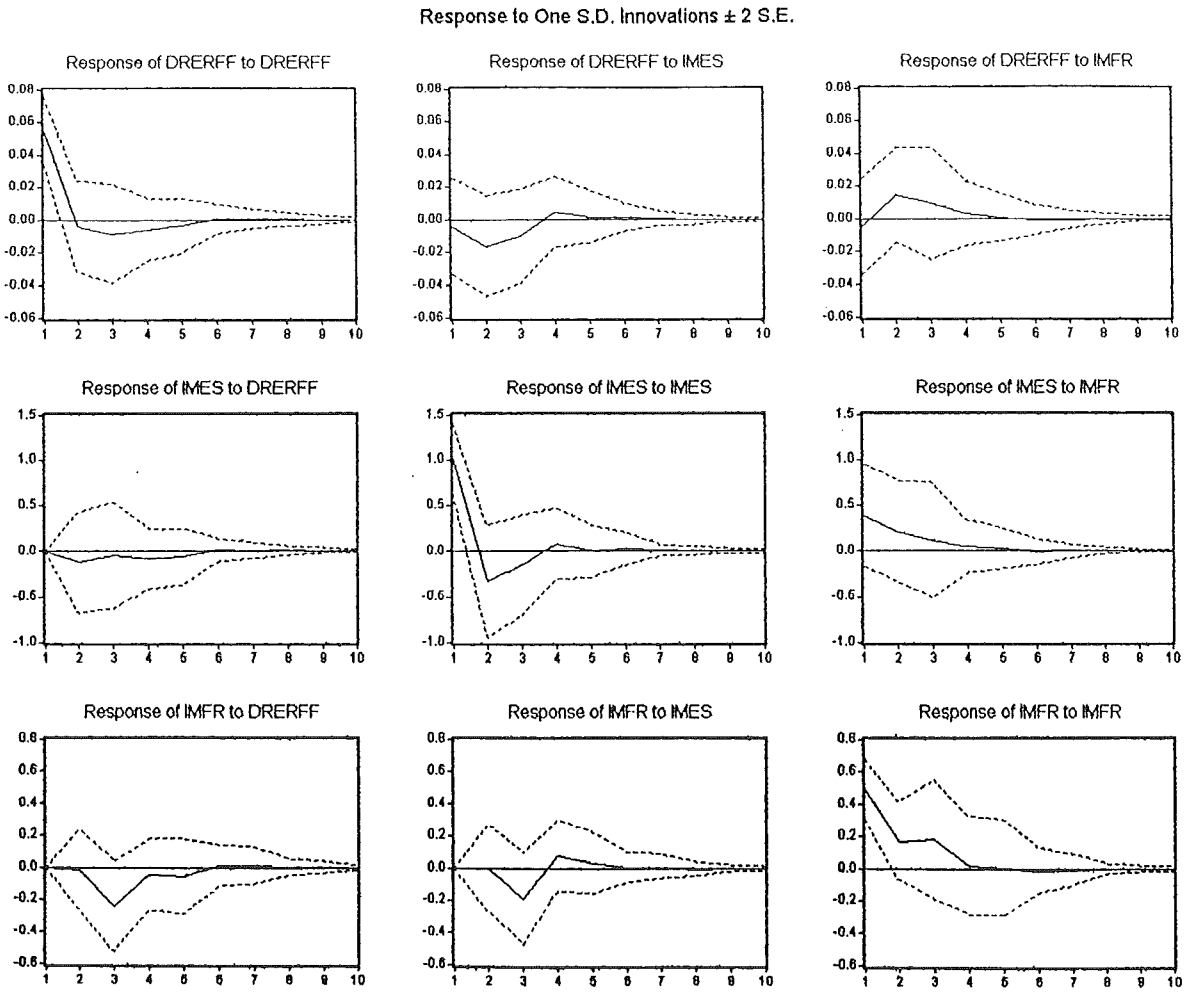
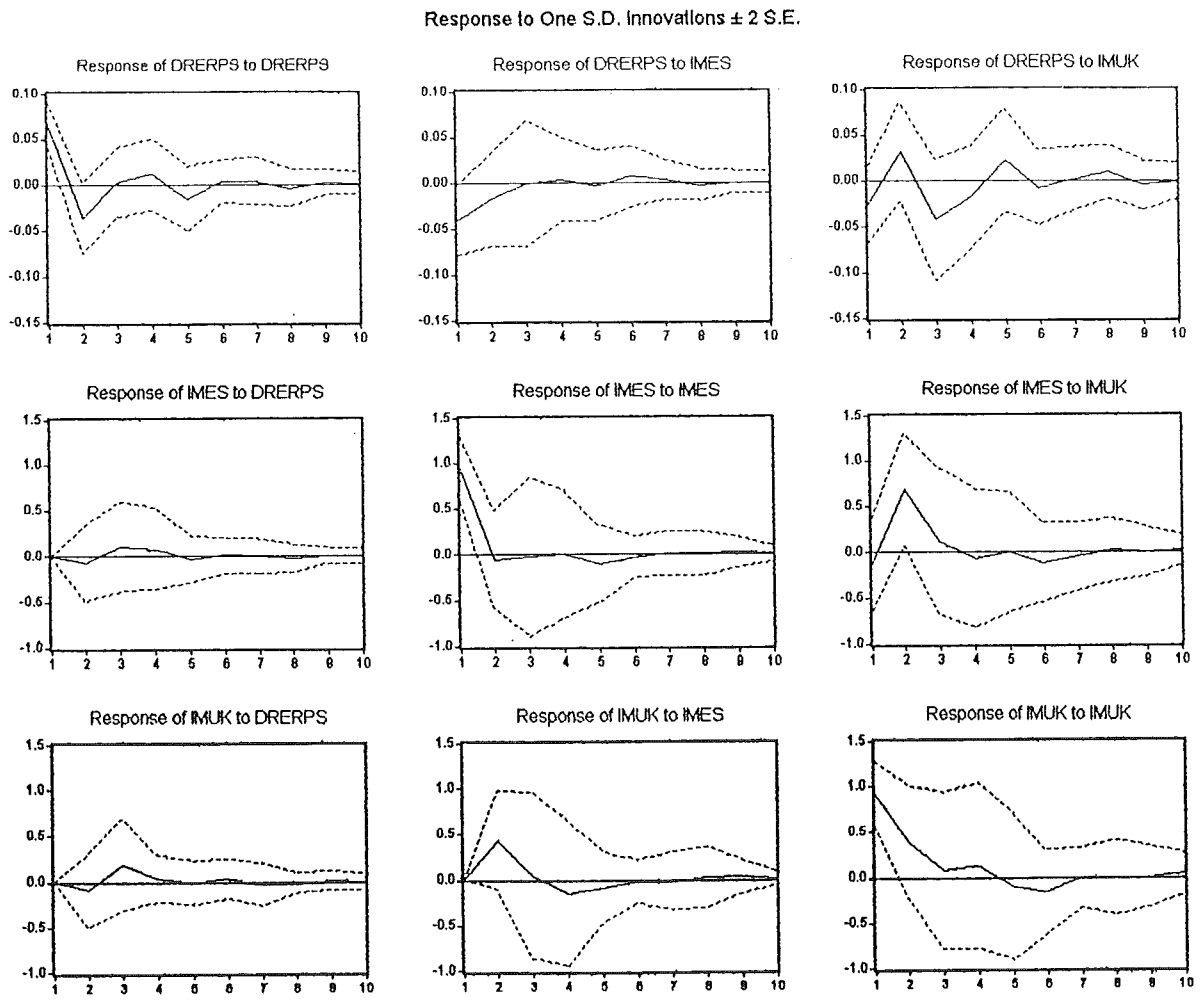


FIGURE 4
IMPULSE-RESPONSE FUNCTIONS
FISCAL IMPULSE MODEL
SPAIN/UNITED KINGDOM
1981-97



respectively. Each of the three rows in the figures, for example, the first one, present the response of the variables (i.e. the first difference of the real exchange rate) to an innovation coming, respectively, from the real exchange rate, the domestic fiscal impulse and the foreign one.

First of all, it should be emphasized that the most important row will be the first one, that is, the case of the real exchange rates (Δrerdm$_t$, Δrerff$_t$, Δrerps$_t$). A devaluation of the currency (that is, an increase) would provoke an initial positive impact for the three currencies. After two periods, the initial depreciation is compensated by appreciations and the effects disappear after 10 periods in the case of the DM and just six for the French franc and the Pound. A positive Spanish fiscal expansion depreciates the pta/DM exchange rate initially, but appreciates the real exchange rates of the other two currencies. Once again, the innovation is compensated quite quickly in the case of the French franc and the Pound Sterling, but takes approximately 10 periods in the Spanish case. Finally, a foreign positive fiscal impulse initially depreciates the real exchange rate although the impact is larger in the German and French models. However, its effects last longer in the case of the pta/DM exchange rate.

Concerning the rest of the variables, the pattern is similar to the one observed for the real exchange rate: the German model is more informative, due to the existence of larger impacts and linkages between the variables. In the other two cases, the response of the variables is only important when it is analyzed the effect of an innovation over the variable itself.

The variance decomposition. The information about the relative importance of the random innovations over every variable is presented in figures 5, 6 and 7 for the German, French and British models, respectively.

First, the upper graphs show the variance decomposition of the first difference of the real exchange rates. This is again the most important case, because it can be used to complement the information about the possible exogeneity of the real exchange rate in the context of the fiscal impulse model. This seems to be the case, because the majority of the variance is explained by its own innovations (80% approximately), with the exception of the British Pound, where the influence of the fiscal impulses is slightly higher. Moreover, as it usually happens with the fiscal policy measures, the impact is not significantly different from zero until the second or third period.

The central graphs show the relative importance of the different shocks on the stance of the Spanish fiscal policy. Apart from its own innovations, which are again the most relevant (over 60%), the fiscal impulses of the foreign countries are more important than the real exchange rates.

FIGURE 5
VARIANCE DECOMPOSITION
FISCAL IMPULSE MODEL
SPAIN/GERMANY
1981-97

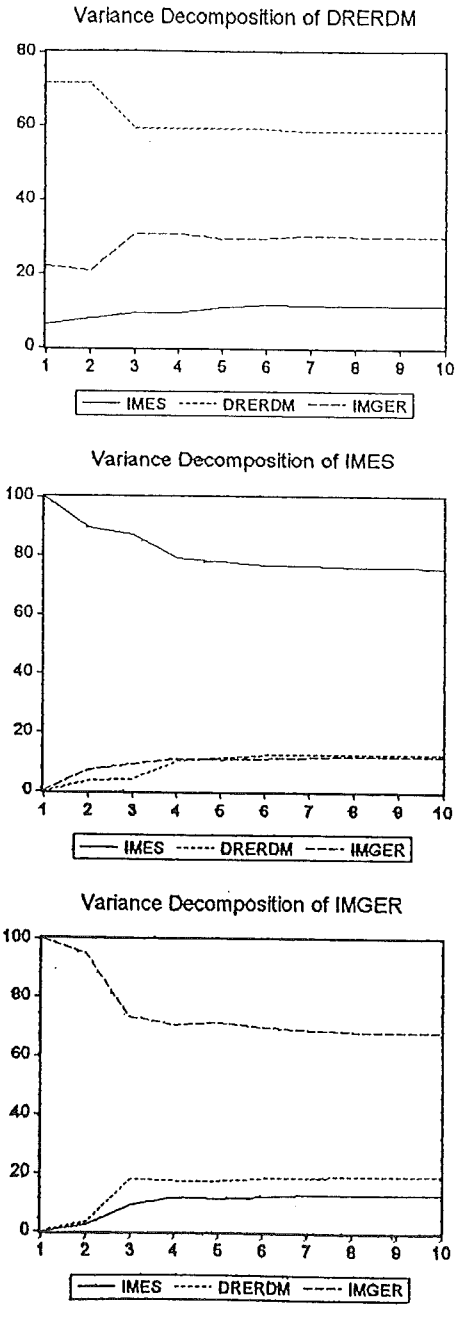


FIGURE 6
VARIANCE DECOMPOSITION
FISCAL IMPULSE MODEL
SPAIN/FRANCE
1981-97

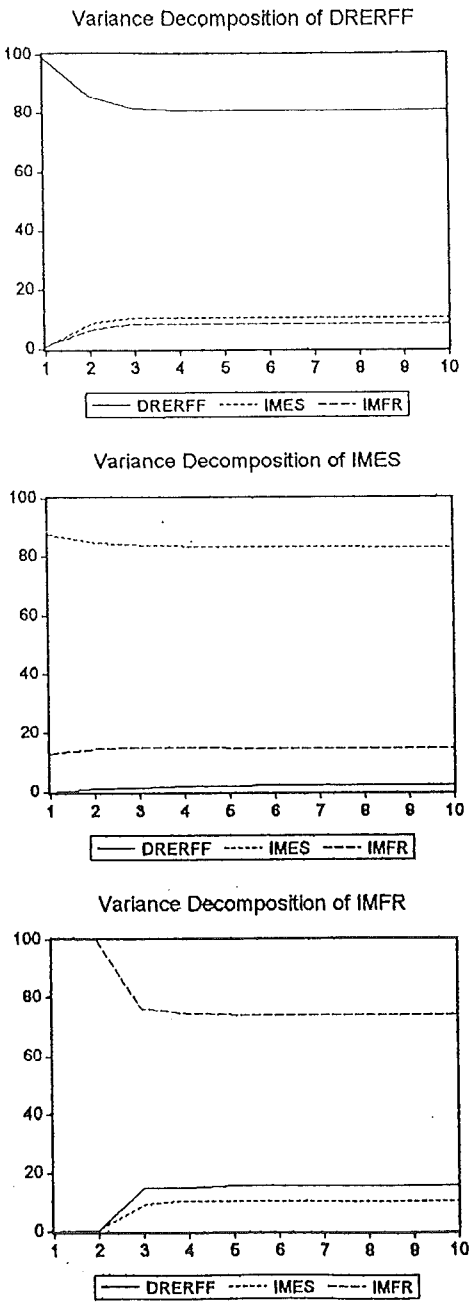
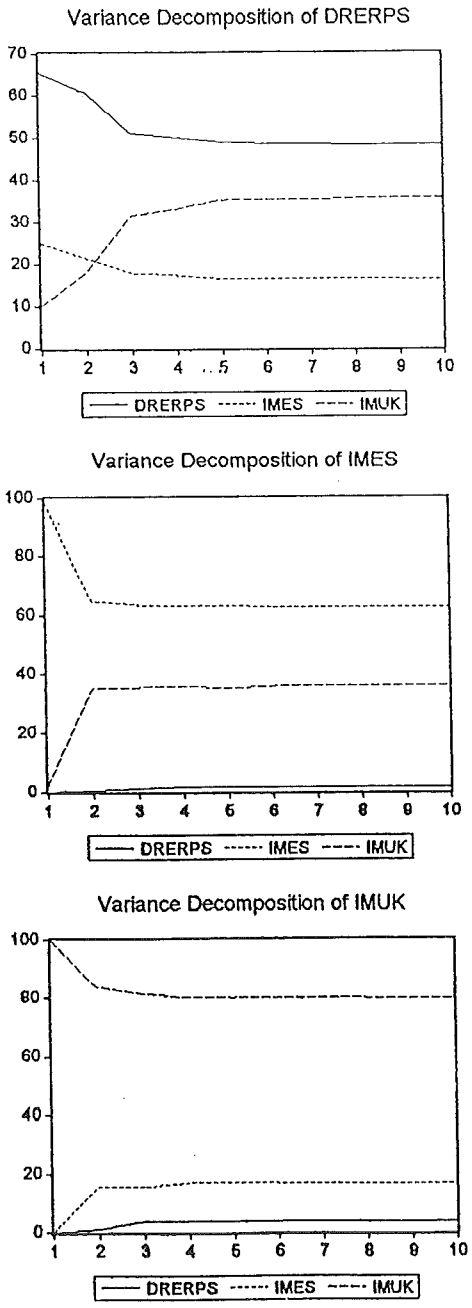


FIGURE 7
VARIANCE DECOMPOSITION
FISCAL IMPULSE MODEL
SPAIN/UNITED KINGDOM
1981-97



Finally, as it is depicted in the lower graphs, the foreign fiscal impulses are more influenced by its own innovations and the real exchange rates than by the Spanish fiscal stance.

Granger causality tests. The bivariate tests for Granger causality are presented in table 4. The null hypothesis is the absence of causality from one variable to the other. The F-tests indicate, in all cases, that the null hypothesis cannot be rejected with very high probabilities. Consequently, there is no causality, in any direction, between the whole set of variables analyzed. This fact confirms the rest of the results, that is, the Obstfeld model is not adequate to explain the bilateral real exchange rates for the samples and countries analyzed.

4.2 Intertemporal real exchange rate model.

From the model developed by Rogoff (1992) and presented in section 3 it is possible to specify an empirical model with the following form:

$$rer_t = \gamma_1 prd_t + \gamma_2 prd_t^* + \gamma_3 g_t + \gamma_4 g_t^* + \gamma_5 rpoil_t \tag{30}$$

where rer_t stands for the real exchange rate, prd_t and prd_t^* are the productivities in the home country and abroad respectively, g_t and g_t^* are the real public expenditure in the domestic and the foreign countries and $rpoil_t$ is the real price of oil. The parameters have the following expected signs: $\gamma_1, \gamma_3, \gamma_5 < 0$, and $\gamma_2, \gamma_4 > 0$. This means that an increase in domestic productivity and public expenditure as well as in the real price of oil would appreciate the domestic currency[14] in real terms, whereas an increase in the foreign variables depreciates it.

Several assumptions have been made in order to derive a testable equation. First, government expenditure should be mostly concentrated in non-tradables, due to the fact that in the model real government spending on tradables has been normalized to zero. A positive instead a negative sign of the domestic variable would mean that the expenditure in tradables is relatively important. Second, it is normally assumed that productivity growth in non-tradables is zero, as in Chinn (1997) or Rogoff (1992). Consequently, productivity in the traded-goods sector is captured using data on output per man per hour, as in Rogoff (1992). Finally, in contrast to the Rogoff

[14] An increase in the real price of oil would appreciate the real exchange rate due to the rise in the costs of the most important input, with a deterioration in the competitive possition of the country.

TABLE 4
GRANGER NON-CAUSALITY TESTS
FISCAL IMPULSE MODEL
1981-97

Null hypothesis	F-Statistic	Probability
$\Delta rerdm_t \nrightarrow imes_t$	0.12	0.88
$imes_t \nrightarrow \Delta rerdm_t$	0.44	0.65
$\Delta rerdm_t \nrightarrow imger_t$	1.05	0.39
$imger_t \nrightarrow \Delta rerdm_t$	1.10	0.37
$\Delta \text{rerff}_t \nrightarrow imes_t$	0.0001	0.99
$imes_t \nrightarrow \Delta \text{rerff}_t$	0.27	0.76
$\Delta \text{rerff}_t \nrightarrow imfr_t$	1.49	0.27
$imfr_t \nrightarrow \Delta \text{rerff}_t$	0.38	0.69
$\Delta rerps_t \nrightarrow imes_t$	0.19	0.82
$imes_t \nrightarrow \Delta rerps_t$	1.86	0.21
$\Delta rerps_t \nrightarrow imuk_t$	0.73	0.50
$imuk_t \nrightarrow \Delta rerps_t$	0.23	0.80

model, it is not assumed the hypothesis of symmetry between the domestic and foreign variables. The series will be considered separately and during the identification process it will be tested whether this restriction is sustained by the data.

4.2.1 Government expenditure, productivity and economic performance in Europe and Spain.

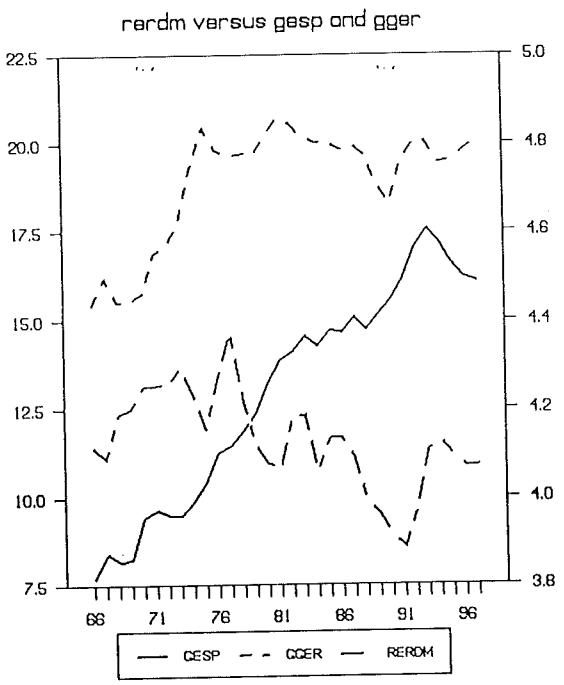
Before starting the analysis of the time series, it should be noted that appendix A describes the data and its sources. The real exchange rate is measured using CPI deflators, which (as Rogoff (1992) points out) includes both the relative price of nontraded goods and the terms of trade, an element that has been excluded from the theoretical model. However, some previous empirical work in Camarero and Tamarit (1998) shows that the real price of oil would capture the majority of the terms of trade shocks.

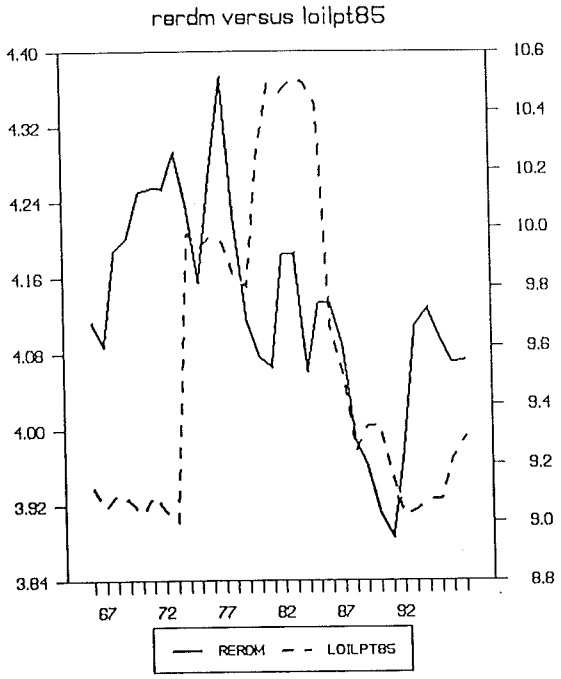
The upper left panel in figures 8, 9 and 10 show the evolution of the real bilateral exchange rates versus the ratio of real government spending to GDP. The three graphs exhibit a negative correlation between the real exchange rates and the domestic fiscal variable, while the correlation is positive with the foreign variable, as predicted by the theory. This fact means that government spending is concentrated in non-traded goods, which was the assumption made by Rogoff (1992). The only case where these correlations are not so evident is the French model (figure 9). As it can be seen in the graph, there are some short episodes where the correlation between the government spending ratio versus the real exchange rate is negative (instead of positive), specially at the beginning of the eighties. This corresponds to the changes introduced during the Mitterand "experiment", when public expenditure grew in France in contrast with the policies implemented in the rest of Europe.

The lower left panel in the three figures plot the real exchange rate (solid line) against the Spanish and foreign country productivities. The productivity variable has been calculated, using the methodology proposed by Martín (1997), as output per man per hour. According to the Rogoff model, the correlation between the real exchange rate and the Spanish productivity must be negative, while it should be positive with the foreign productivity. The British case is the only one where this relation is quite clear. Some doubts can be raised in the other two countries, due to the general upward trend in productivity. This fact supports the Rogoff (1992) model, which predicts that productivity may have little explanatory power, at least in the short-run.

Finally, the upper right panels of figures 8, 9 and 10 plot the three real exchange rates of the peseta, respectively, against the real price of oil. There is a negative correlation between these two variables both in the German

FIGURE 8
REAL EXCHANGE RATES, GOVERNMENT SPENDING, PRODUCTIVITY AND REAL PRICE OF OIL
SPAIN AND GERMANY





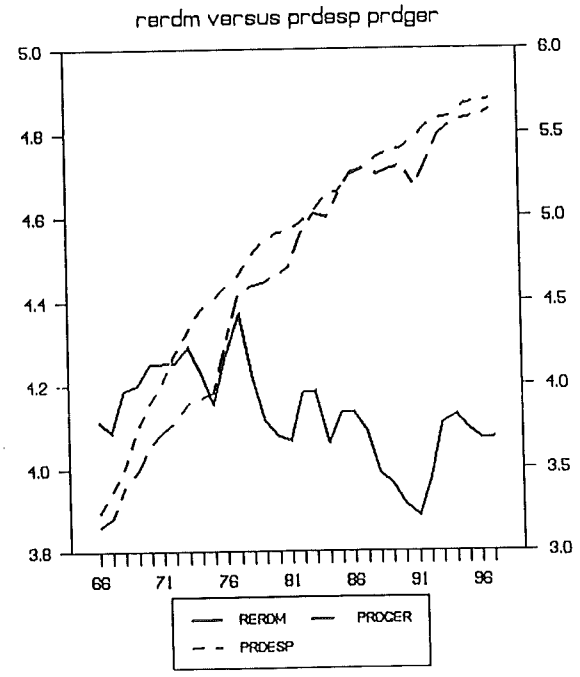


FIGURE 9
REAL EXCHANGE RATE, GOVERNMENT SPENDING, PRODUCTIVITY AND REAL PRICE OF OIL
SPAIN AND FRANCE

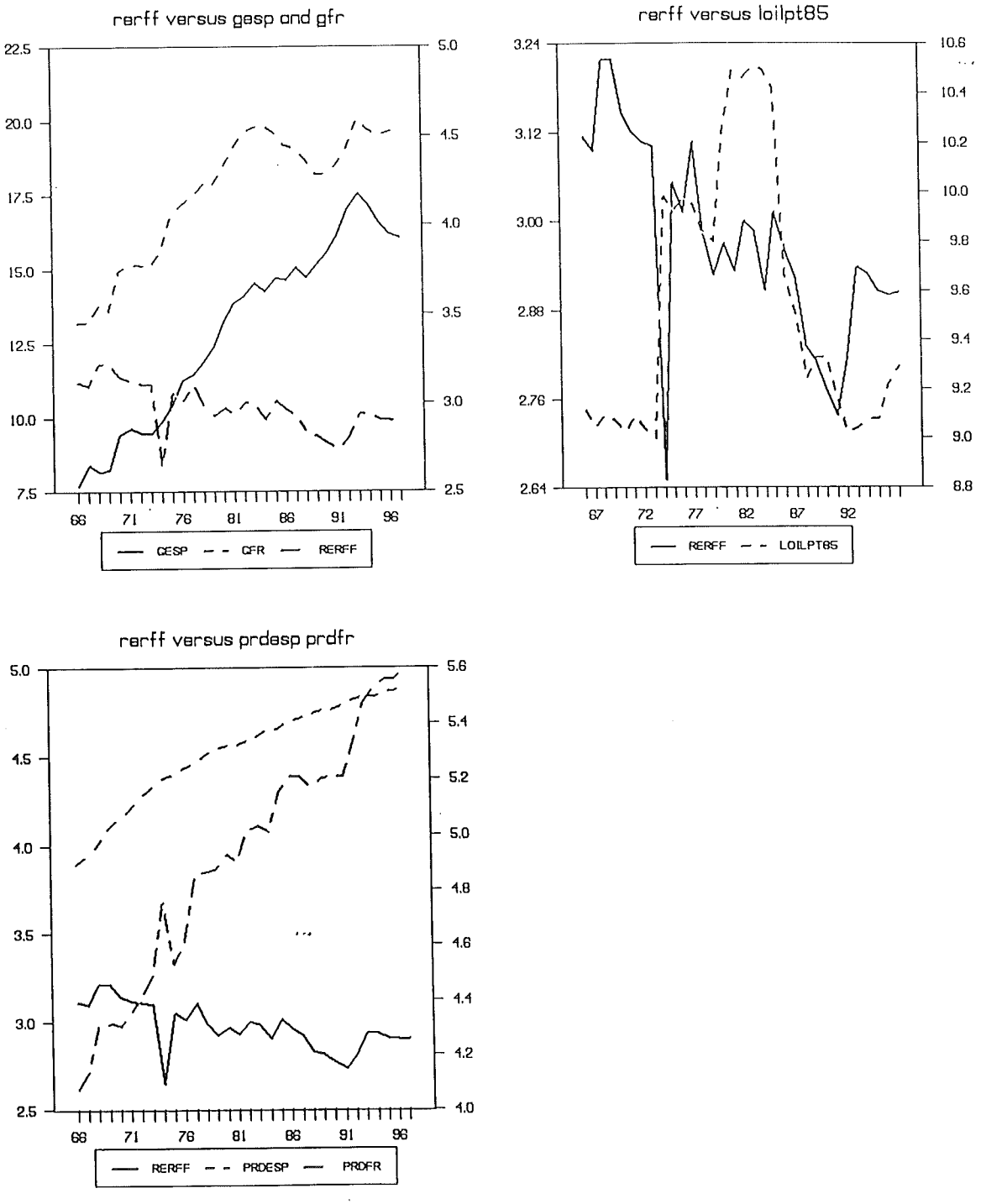
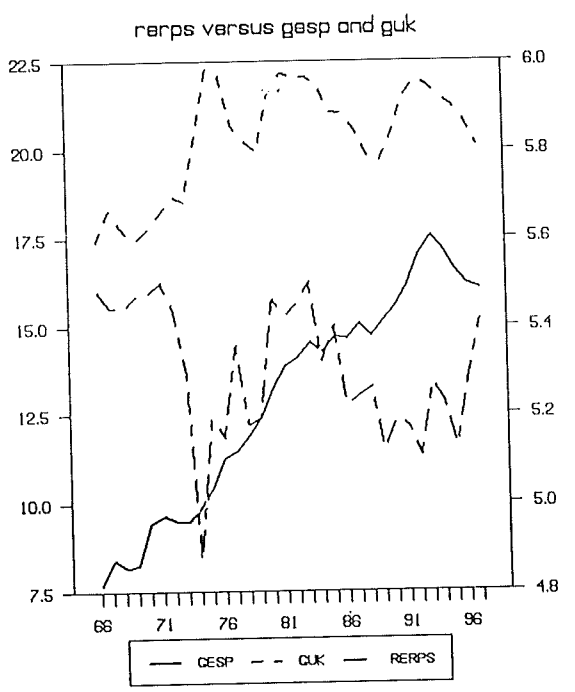
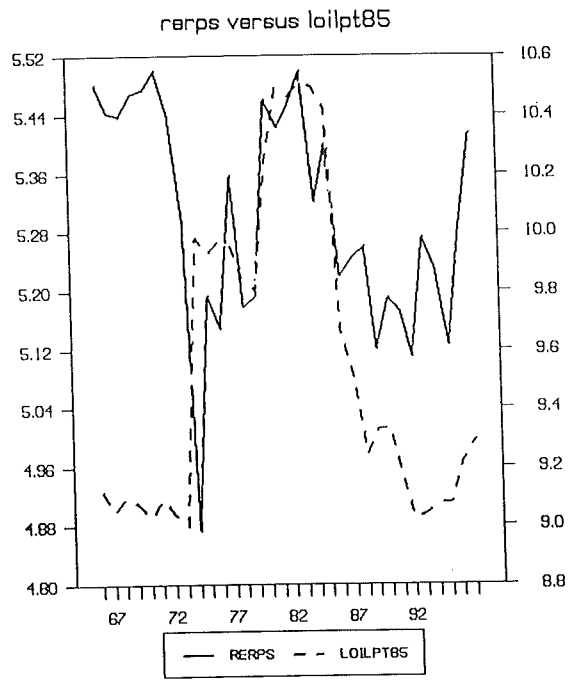
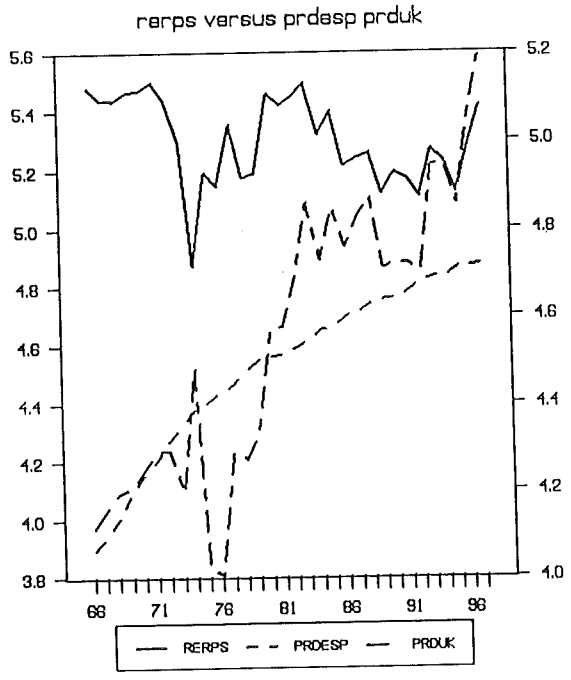


FIGURE 10
REAL EXCHANGE RATE, GOVERNMENT SPENDING PRODUCTIVITY AND REAL PRICE OF OIL
SPAIN AND GREAT BRITAIN







(figure 8) and the French model (figure 9), due to the dependence on oil existing in the three countries involved. The picture is different in the British case: since the seventies, when the UK started extracting oil, the correlation is inverse, showing a positive relation, except for some years at the end of the sample. Consequently, this variable (or even the model) may not be appropriate for the British economy.

4.2.2 Order of integration of the variables.

The univariate analysis of the series is presented in table 5. The first part of the table shows the results of the Phillips-Perron tests for both the series in differences and in levels, where the null hypothesis is the existence of a unit root. According to the Phillips-Perron tests, it is possible to reject the existence of two unit roots in all the variables analyzed. The right-hand part of the table presents the results for the variables in levels. In general, it is not possible to reject the existence of a unit root in the series. There are, however some doubts concerning two productivity variables ($prdesp_t$, $prdfr_t$) and two real exchange rates ($rerff_t$, $rerps_t$).

The KPSS tests may help to decide the order of integration of the more borderline series. The results are presented in the second part of table 5. In this case, the null hypothesis is stationarity, which cannot be rejected for all the variables in differences except for $prdesp_t$. For the levels of the series, it is possible to reject stationarity in all the cases analyzed.

Consequently, the main conclusion is that all the variables contain a unit root, although, due to the relatively long span of time of the sample (1966-97), some variables may contain a structural break. This is, precisely, the case of productivity and real exchange rates, where technological changes or devaluations may affect the power of the tests to detect a unit root.

4.2.3 Cointegration results.

This subsection is devoted to the estimation of the empirical model for three bilateral real exchange rates of the peseta: the Pta/DM, Pta/French frank and Pta/Pound Sterling.

Bilateral model between Spain and Germany. Once we have concluded in the univariate analysis that the variables are integrated of order one, it is possible to proceed and estimate the cointegrating relations using the Johansen procedure. The sample period goes from 1966 to 1997 and the results are presented in table 6. According to the $\lambda-\max$ and trace tests and taking into account the eigenvalues (note that the critical values in the table

TABLE 5
UNIT ROOT TESTS
ROGOFF INTERTEMPORAL MODEL
1966-1997

Variables	Phillips-Perron tests			Phillips-Perron tests		
	$Z(t_{\tilde{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$	$Z(t_{\tilde{\alpha}})$	$Z(t_{\alpha*})$	$Z(t_{\hat{\alpha}})$
	Series in differences, $l = 1$			Series in levels, $l = 1$		
$gesp_t$	-4.42***	-4.21***	-3.46***	-0.88	-1.66	2.50
$gger_t$	-4.49***	-4.41***	-4.27***	-1.77	-2.15	1.07
gfr_t	-4.86***	-4.34***	-3.69***	-1.26	-2.08	1.97
guk_t	-3.93**	-3.84***	-3.85***	-1.94	-2.25	0.37
$prdesp_t$	-5.14**	-2.51	-1.59	-2.74	-7.71***	5.80
$prdger_t$	-5.06***	-4.51***	-2.94***	-1.31	-2.03	3.83
$prdfr_t$	-7.23***	-7.13***	-5.21***	-3.86**	-1.18	3.39
$prduk_t$	-7.80***	-7.73***	-7.12***	-3.15	-0.62	1.57
$rerdm_t$	-4.87***	-4.84***	-4.84***	-2.92	-2.03	0.14
$rerff_t$	-8.86***	-8.88***	-8.84***	-4.44***	-2.86*	-0.49
$rerps_t$	-7.65***	-7.54**	-39.15***	-3.18	-3.12**	-0.02
$loilpt85_t$	-5.09***	-4.94***	-4.94***	-1.51	-1.53	0.03

The critical values for the Phillips-Perron tests are the following:

Signif. Level	$Z(t_{\tilde{\alpha}})$	$Z(t_{\alpha^*})$	$Z(t_{\hat{\alpha}})$
1%	-4.38	-3.75	-2.66
5%	-3.60	-3.00	-1.95
10%	-3.24	-2.63	-1.60

in table 4, (*), (**) and (***) denote, respectively, rejection of the null hypothesis of a unit root at 10, 5 and 1% significance levels. The critical values for $Z(t_{\tilde{\alpha}})$, $Z(t_{\alpha^*})$ and $Z(t_{\hat{\alpha}})$ are taken from Fuller (1976), table 8.5.2.

TABLE 5 (Continued)
UNIT ROOT TESTS
ROGOFF INTERTEMPORAL MODEL
1966-1997

Variables	KPSS tests, $l = 1$		KPSS tests, $l = 1$	
	η_μ	η_τ	η_μ	η_τ
	Series in differences		Series in levels	
$gesp_t$	0.26	0.08	1.62***	0.21***
$gger_t$	0.23	0.06	0.92***	0.33***
gfr_t	0.32	0.06	1.37***	0.35***
guk_t	0.16	0.03	0.81***	0.24***
$prdesp_t$	1.15***	0.15**	1.58***	0.37***
$prdger_t$	0.31	0.03	1.62***	0.36***
$prdfr_t$	0.07	0.03	1.63***	0.19**
$prduk_t$	0.07	0.05	1.48***	0.10
$rerdm_t$	0.06	0.05	0.85***	0.12*
$rerff_t$	0.04	0.03	1.03***	0.08
$rerps_t$	0.09	0.05	0.27*	0.08
$loilpt85_t$	0.18	0.08	0.34*	0.34***

The critical values for the KPSS tests are the following:

Signif. Level	η_μ	η_τ
1%	0.73	0.21
5%	0.46	0.14
10%	0.34	0.11

in table 4, (*), (**) and (***) denote, respectively, rejection of the null hypothesis of stationarity at 10, 5 and 1% significance levels. The critical values for η_μ and η_τ have been taken from Kwiatkovski, Phillips, Schmith and Shin (1992), table 1.

are displayed at 90% significance level), there would be approximately 3 or 4 cointegrating vectors. This can be also complemented with the information contained in the roots of the companion matrix, that has three roots equal to unity[15]. Consequently, the final choice is three cointegrating vectors.

The unrestricted model together with the identified (restricted) model are presented in the second part of table 6. Following the procedures described in Johansen (1995) and Hansen and Juselius (1995), the model has been completely identified obtaining a relationship between the real exchange rate and foreign and domestic productivity (vector β_1'), the Rogoff model (vector β_2') and a relationship between the real exchange rate and the domestic and foreign real public spending relative to GDP (vector β_3'). These restrictions have been accepted using LR tests with a 6% probability (that is, larger than the 5% required). Previous to the estimation of the complete model, it has been tested the existence of a stationary relationship between the real exchange rate and the pairs of variables, which was very helpful and informative at the time of the identification process.

However, we should concentrate now on vector β_2' which is the one of interest for the purpose of this paper. In contrast to the results found by Chinn (1997), who could not estimate a model displaying the correct signs and magnitudes of the parameters involved, vector β_2' satisfied the restrictions implied by the Rogoff (1992) model. First, the domestic and the foreign variables have the same magnitude but opposite sign and those signs are the expected according to the theoretical model. Not only the productivity coefficients but also the government spending ones are smaller than unity and the sign of the real price of oil is negative, as predicted by the model. Moreover, this last variable is weakly exogenous (with a probability of 30%) and only enters the Rogoff equation.

Thus, the model developed by Rogoff is supported by the data for the case of the peseta/DM real exchange rate.

Bilateral model between Spain and France. The second bilateral exchange rate model analyzed relates the Peseta and the French Franc. The results, for I(1) variables, are presented in table 7. The $\lambda - \max$ and the trace test recomend, again in this case, to choose a number of cointegrating vectors between 3 and 4. Once more, the use of the information contained in the roots of the companion matrix suggests a choice of just three vectors.

The next step consists of estimating the model and, afterwards, identifying it. Both the restricted and unrestricted version appear in the lower

[15]This information is not included in the paper but can be provided by the author upon request.

TABLE 6
JOHANSEN COINTEGRATION TESTS
PTA/DM
1966-97

Eigenvalues	$\lambda -$ max test	Trace test	$H_0 : r$	$p - r$	$\lambda -$ max 90%	Trace 90%
0.8683	60.81*	188.00*	0	6	24.63	89.37
0.7938	47.37*	127.18*	1	5	20.90	64.74
0.6771	33.91*	79.81*	2	4	17.15	43.84
0.5859	26.45*	45.90*	3	3	13.39	26.70
0.3705	13.88*	19.46*	4	2	10.60	13.31
0.1695	5.57*	5.57*	5	1	2.71	2.71

Variables	Unrestricted model			Restricted model		
	β_1	β_2	β_3	β_1	β_2	β_3
$rerdm_t$	-3.841	1	1.029	8.774	1	-32.024
$prdesp_t$	0.529	-1.539	1	1	0.498	0
$prdger_t$	-6.178	0.459	-0.254	-2.276	-0.498	0
$gesp_t$	1	0.005	-0.272	0	0.009	1
$gger_t$	0.217	0.061	0.511	0	-0.009	0.055
$loilpt85_t$	-0.264	-0.134	-0.264	0	0.026	0

Restrictions: LR tests		
Type of test	χ^2	Probability value
Identifying restrictions	$\chi^2(2) : 5.58$	0.06
Weak exogeneity of $loilpt85_t$	$\chi^2(5) : 6.06$	0.30

NOTE: the critical values for the $\lambda -$ max and *trace* tests are taken from Osterwald-Lenum (1992).

TABLE 7
JOHANSEN COINTEGRATION TESTS
PTA/FF
1966-97

Eigenvalues	$\lambda - \max$ test	Trace test	$H_0 : r$	$p - r$	$\lambda - \max$ 90%	Trace 90%
0.8509	57.09*	145.88*	0	6	24.63	89.37
0.6555	31.97*	88.79*	1	5	20.90	64.74
0.6265	29.54*	56.82*	2	4	17.15	43.84
0.4282	16.77*	27.28*	3	3	13.39	26.70
0.2128	7.18	10.51	4	2	10.60	13.31
0.1051	3.33	3.33	5	1	2.71	2.71

Variables	Unrestricted model			Restricted model		
	β_1	β_2	β_3	β_1	β_2	β_3
$rerff_t$	0.450	1	1	-1.973	1	1
$prdesp_t$	1	1.923	0.567	1	0	0.075
$prdfr_t$	0.077	-0.818	-0.110	-1.675	0	-0.075
$gesp_t$	0.033	0.098	-0.017	0	0.043	0.047
gfr_t	-0.162	-0.240	0.036	0	-0.039	-0.047
$loilpt85_t$	0.232	0-276	-0.098	0	0	0.006
Restrictions: LR tests						
Type of test				χ^2		Probability value
Identifying restrictions				$\chi^2(2) : 6.3$		0.05
Weak exogeneity of $loilpt85_t$				$\chi^2(5) : 6.88$		0.23

NOTE: the critical values for the $\lambda - \max$ and *trace* tests are taken from Osterwald-Lenum (1992).

part of table 7, as well as the LR tests for identification and weak exogeneity of $loilpt85_t$. The three restricted vectors contain the relationship between the real exchange rate and the two productivities (vector β_1'), the relation between the real exchange rate and the two public expenditure variables (vector β_2') and, finally, the Rogoff model (vector β_3'). These restrictions are acceptable with a probability of 5%.

Vector β_3' in the French model is very similar to the one estimated for the German one. The signs are again "correct" according to the theoretical predictions, and the coefficients are smaller than one in absolute value, being acceptable the restriction of symmetry between domestic and foreign variables. However, their magnitudes are different, if compared to the German case: smaller for the productivity variables and larger for the fiscal ones. In any case, the French model also satisfies the restrictions of the theoretical model of Rogoff (1992).

Bilateral model between Spain and the United Kingdom. The last model that has been analyzed relates Spanish and British variables. However, the study of the time series properties of the variables involved (see table 8) did not recommend to continue with the implementation of the Johansen procedure. In fact, the bilateral real exchange rate of the Peseta versus the Pound Sterling turned out to be weakly exogenous relative to the rest of the variables of the model. This means that this real exchange rate is not determined by the variables of the system specified according to the Rogoff model. It should be borne in mind that the British economy has particular features that makes its cycle differ from the so-called "continental" cycle. The lower degree of integration and the asymmetries in the evolution of Britain compared to the rest of Europe may explain the absence of a relation between the British and the Spanish variables in order to explain the real exchange rate in the context of the Rogoff (1992) model.

5 Conclusions.

In this paper, we apply two types of models to the determination of the real bilateral exchange rate of the peseta against the DM, French franc and Pound Sterling. The first approach is based on a Keynesian asset model for the determination of the real exchange rate with rational expectations proposed by Obstfeld (1985) and has been previously applied to the American dollar by Chinn (1997). The solution of the model gives rise to an expression for the real exchange rate that depends on real aggregate demand shocks which are originated in fiscal domestic and foreign variables. The variables that

TABLE 8
ROGOFF MODEL
TIME SERIES PROPERTIES
PTA/PS
1966-97

Test for lon-run exclusion								
Coint. relations	DGF	$\chi^2(*)$ at 5%	$rerps_t$	$prdesp_t$	$prduk_t$	$gesp_t$	guk_t	$loilpt85_t$
1	1	3.84	10.64*	1.11	0.40	3.35	17.28*	20.27*
2	2	5.99	14.98*	4.39	0.46	5.33	23.27*	24.27*
3	3	7.81	19.72*	9.26*	0.57	7.04	24.16*	27.11*
4	4	9.49	26.71*	12.21*	6.88	9.16	29.61*	30.83*
5	5	11.07	27.58*	12.72*	7.66	9.69	29.81*	30.83*
Test for stationarity								
Coint. relations	DGF	$\chi^2(*)$ at 5%	$rerps_t$	$prdesp_t$	$prduk_t$	$gesp_t$	guk_t	$loilpt85_t$
1	5	11.07	44.84*	29.04*	40.79*	40.43*	36.42*	48.47*
2	4	9.49	18.56*	8.06	12.87*	13.09*	3.84	16.09*
3	3	7.81	11.66*	8.04*	7.29	11.61*	0.03	9.81*
4	2	5.99	8.05*	6.30*	2.84	6.18*	0.01	7.43*
5	1	3.84	0.82	0.19	0.34	0.00	0.00	0.32
Test for weak exogeneity								
Coint. relations	DGF	$\chi^2(*)$ at 5%	$rerps_t$	$prdesp_t$	$prduk_t$	$gesp_t$	guk_t	$loilpt85_t$
1	1	3.84	2.65	11.18*	0.73	15.23*	4.46*	0.37
2	2	5.99	3.04	18.00*	2.67	15.24*	5.63	0.88
3	3	7.81	7.66	21.59*	2.68	16.37*	9.15*	1.75
4	4	9.49	7.73	22.45*	6.60	21.55*	15.68*	4.68
5	5	11.07	8.64	23.01*	7.51	22.46*	16.26*	4.71

Obstfeld (1985) himself proposed in order to test this model are the *fiscal impulses* as defined by the IMF.

The second approach is intertemporal and based on a fixed-factor model derived by Rogoff (1992). The most interesting feature of this model is the fact that the capital market is assumed to be in a process of liberalization and the factors are sector-specific. Consequently, in the short and medium run and, in contrast to the Balassa-Samuelson type models, the productivity differential could not be very explanatory. Thus, in this context, the demand-side, as represented by public spending, would be the main determinant of real exchange rates. However, in the long-run, all the factors would be mobile and both demand and supply variables would determine the real exchange rate.

The cointegration theory has been the econometric method selected to test and, if possible, compare both models. However, it should be emphasized that the data and the results are not directly comparable. The IMF fiscal impulses are only available for the Spanish economy for the sample period 1981-1997. As a consequence, the short span of data obliges us to be careful at the time of interpreting the empirical results. Although Chinn (1997) applied a MA filter to obtain quarterly data for the IMF series, Campbell and Perron (1991), for example, did not recommend such transformations, due to the lack of additional information content provided by this kind of transformations. The unit root test pointed out at the stationarity of the fiscal impulses, a fact that should not surprise the applied researcher if one examines the definition of the variables themselves, as it is shown in section 2.2. Another discouraging result related to the ability of this model to explain the real exchange rate is that this variable seems to be weakly exogenous for the three cases considered: that is, the fiscal impulses did not seem to be very explanatory of the real exchange rates. However, even in the absence of cointegration, it is still possible to study the effects of the fiscal impulses shocks on the real exchange rate using the VAR methodology. The results are in general not very informative, except for the case of the impulse-response functions for the German case. Finally, the lack of support of this model by the data is confirmed by the Granger non-causality tests that did not allow us to accept causation from any of the variables involved in the model.

The intertemporal model proposed by Rogoff (1992) has been also tested using the cointegration methodology, although, in this case, the sample period is longer, going from 1966 to 1997. The solution of the model relates the real exchange rate with domestic and foreign productivity and public spending ratios to GDP, as well as the real oil prices. The variables are all first order integrated and it has been possible to estimate and identify the models corresponding to the pta/DM and the pta/FF cases. However, the

British model has not been estimated due to the weak exogeneity of the real exchange rate of the peseta against the Pound Sterling in relation to the rest of the variables in Rogoff (1992). The identification of the models for the other two exchange rates give rise to similar equations that relate the real exchange rate with the whole set of variables, as well as to the productivities and the fiscal expenditure variables separately. Moreover, it is accepted the symmetry restriction (both for the German and French case) in the equation representing the Rogoff (1992) model, a fact that Rogoff assumed and did not test. Consequently, the sample and data analyzed support the intertemporal approach in contrast to the Keynesian model, although future research would be necessary in order to determine the relative importance of demand and supply variables in this model.

A Sources and data.

f_t : fiscal impulses, IMF World Economic Outlook. Data kindly provided by Randa Sab.

fes_t : Spanish fiscal impulses.

$fger_t$: German fiscal impulses.

$ffra_t$: French fiscal impulses.

fuk_t : British fiscal impulses.

rer_t : logarithm of the bilateral real exchange rate of the Spanish peseta calculated using consumer prices. Source: Cronos.

$rerdm_t$: real bilateral exchange rate peseta/Deutche Mark.

$rerff_t$: real bilateral exchange rate peseta/French Franc.

$rerps_t$: real bilateral exchange rate peseta/Pound Sterling.

g_t : real government expenditure relative to real GDP. Basis: 1985. Source: Cronos.

$gesp_t$: real government expenditure to real GDP in Spain.

$gger_t$: real government expenditure to real GDP in Germany.

gfr_t : real government expenditure to real GDP in France.

guk_t : real government expenditure to real GDP in the United Kingdom.

prd_t : productivity per man per hour,that has been calculated using the number of workers in the different countries and assuming that the weekly hours are 40. Source: Cronos for employment and real output.

$prdesp_t$: Spanish productivity.

$prdger_t$: German productivity.

$prdfr_t$: French productivity.

$prduk_t$: British productivity.

$loilpt85_t$: logarithm of the real price of oil in Spanish pesetas using CPI base 1985. Source: Ministerio de Economía y Hacienda. Dirección General de Previsión y Coyuntura.

B Some concepts and methodological aspects.

B.1 Definition of weak exogeneity in the Johansen (1988) model.

According the Johansen (1988), a p-dimensional time series vector X_t admitis a VAR representation:

$$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + ... + \Pi_k X_{t-k} + \mu + \Phi D_t + \varepsilon_t \qquad (31)$$

where $\varepsilon_1, ..., \varepsilon_T$ are $IIN(0, \Lambda)$, $X_0, X_{-1}, ..., X_{-k+1}$ are fixed, μ is a vector of constants and D_t are centered seasonal dummies. If the matrix polynomial, $\Pi(z)$ is such that $\mid \Pi(1) \mid = 0$, the long run impact matrix:

$$\Pi = I - \Pi_1 - \Pi_2 - ... - \Pi_k$$

has reduced rank, $r < p$, where p is the number of variables in the system. This matrix can be written as the product of two $(p \times r)$ matrices, $\Pi = \alpha\beta'$. In short, X_t is not stationary as a vector process, though we will assume that ΔX_t and the r linear combinations $\beta' X_t$ are stationary. Thus, the r columns of β are the cointegrating vectors.

The ECM representation of the Johansen (1988) model can be expressed in the following form:

$$\Delta X_t = \alpha\beta' X_{t-1} + \Gamma_1 \Delta X_{t-1} + ... + \Gamma_{k-1} \Delta X_{t-k+1} + \mu + \Phi D_t + \varepsilon_t \qquad (32)$$

where $\Gamma_i = -(\Pi_{i+1} + ... + \Pi_k)$. It should be noted that a VAR in first differences would suffer from misspecification unless $\Pi = 0$, whereas specification of a VAR in levels would not take into account the cross equation restrictions implied by the hypothesis of cointegration, $\Pi = \alpha\beta'$. The short-run dynamics of the model are determined by the ECM. It specifies, according to Gardeazábal and Regúlez (1992), how the variables change over time as a function of four components: the deviations to the r long-run equilibrium relationships; past changes in al variables; a purely deterministic component; and a stochastic disturbance.

The first component is the most important one: the loadings matrix α, measures how the different variables in the system react to the stationary equilibrium errors $\beta' X_t = Y_t$. Thus, a generic element of this matrix, α_{ij},

measures the force of adjustment with which the $i-th$ variable of the system reacts to deviations from the $j-th$ long-run relationship. Then, if $\alpha_{ij}=0$ for $j=1,...,r$, the $i-th$ variable can be considered weakly exogenous with respect to the parameters of interest β, so that the estimation of β can be performed conditional on the $i-th$ variable, thus , reducing by one the dimension of the system.

B.2 Stationary VAR analysis.

A group of time series can be written in terms of an autoregression:

$$y_t = c + \phi_1 y_{t-1} + \phi_2 y_{t-2} + ... + \phi_p y_{t-p} + \varepsilon_t \tag{33}$$

where

$$E(\varepsilon_t) = 0$$

$$E(\varepsilon_t \varepsilon_\tau) = \sigma^2 \text{ for } t = \tau \text{ and } 0 \text{ otherwise}$$

The same VAR model can be expressed in MA(∞) form (see Hamilton, 1994):

$$y_t = \mu + \varepsilon_t + \Psi_1 \varepsilon_{t-1} + \Psi_2 \varepsilon_{t-2} + ... \tag{34}$$

Thus, the matrix Ψ_s can be interpreted as:

$$\frac{\partial y_{t+s}}{\partial \varepsilon_t'} = \Psi_s \tag{35}$$

that is, the row i column j element Ψ_s identifies the consequences of a one-unit increase in the jth variable's innovation at date t (ε_{jt}) for the value of the ith variable at time $t+s$ ($y_{i,t+s}$), holding all other innovations at all dates constant.

A plot of the row i, column j element of Ψ_s,

$$\frac{\partial y_{i,t+s}}{\partial \varepsilon_{jt}} \tag{36}$$

as a function of s is called the *impulse-response function.* It describes the response of $y_{i,t+w}$ to a one-time impulse in y_{jt} with all other variables dated t or earlier held constant.

The ambiguity in interpreting impulse-response functions arises from the fact that the errors are never totally uncorrelated and the effect of any common component (which cannot be identified with any specific variable) is normally attributed to the variable that comes first. The errors are orthogonalized by a Cholesky decomposition so that the covariance matrix of the resulting innovations is diagonal. While the Cholesky decomposition is widely used, it is a rather arbitrary method of attributing common effects[16].

The error in forecasting the VAR s periods ahead can be written as:

$$y_{t+s} - \hat{y}_{t+s|t} = \varepsilon_{t+s} + \Psi_1 \varepsilon_{t+s-1} + \Psi_2 \varepsilon_{t+s-2} + ... + \Psi_{s-1} \varepsilon_{t+1} \tag{37}$$

Then, the mean squared error of this s-period-ahead forecast is:

$$\begin{aligned} MSE(\hat{y}_{t+s|t}) &= E[(y_{t+s} - \hat{y}_{t+s|t})(y_{t+s} - \hat{y}_{t+s|t})'] \\ &= \Omega + \Psi_1 \Omega \Psi_1' + \Psi_2 \Omega \Psi_2' + \Psi_{s-1} \Omega \Psi_{s-1}' \end{aligned} \tag{38}$$

where

$$\Omega = E(\varepsilon_t \varepsilon_t')$$

It is possible to calculate the contribution of each of the orthogonalized disturbances $(u_{1t}, ..., u_{nt})$ to the MSE. Through some algebra, the matrix Ω can be written as a function of the variance of the errors (that are uncorrelated). Substituting, the MSE of the s-period-ahead forecast can be written as the sum of n terms, one arising from each of the disturbances u_{jt} :

$$\begin{aligned} MSE(\hat{y}_{t+s|t}) = \sum_{j=1}^{n} \{ & Var(u_{jt}) \cdot [a_j a_j' + \Psi_1 a_j a_j' \Psi_1' + \\ & + \Psi_2 a_j a_j' \Psi_2' + ... + \Psi_{s-1} a_j a_j' \Psi_{s-1}'] \} \end{aligned} \tag{39}$$

where a_j is the jth element of matrix A (which is the lower triangular matrix associated to Ω with 1s along the principal diagonal). With expression (39), it is possible to calculate the contribution of the jth orthogonalized innovation to the MSE of the s-period-ahead forecast:

[16]Note that the ordering of the equations can change the results. For a wide discussion of this issue see Hamilton (1994).

$$Var(u_{jt}) \cdot [a_j a_j' + \Psi_1 a_j a_j' \Psi_1' + \Psi_2 a_j a_j' \Psi_2' + ... + \Psi_{s-1} a_j a_j' \Psi_{s-1}'] \qquad (40)$$

This magnitude depends on the ordering of the variables, as in the impulse-response function. As $s \to \infty$ for a covariance-stationary VAR, $MSE(\hat{y}_{t+s|t}) \to \Gamma_0$, the unconditional variance of the vector y_t. Thus, the equation (39) permits calculation of the portion of the total variance of y_i that is due to the disturbance u_j by letting s become suitably large. This defines the *variance decomposition.*

Finally, one of the key questions related to vector autoregressions is how useful some variables are for forecasting others. A common test for this issue was proposed by Granger (1969) and popularized by Sims (1972). In the case of *bivariate Granger causality* it is investigated whether a scalar y can help forecast another scalar x. If it cannot, then we say that y does not Granger-cause x. Formally, *y fails to Granger-cause x* if for all $s > 0$ the mean squared error of a forecast of x_{t+s} based on $(x_t, x_{t-1}, ...)$ is the same as the MSE of a forecast of x_{t+s} that uses both $(x_t, x_{t-1}, ...)$ and $(y_t, y_{t-1}, ...)$. If we restrict ourselves to linear functions, y fails to Granger-cause x if:

$$\begin{aligned} &MSE[\hat{E}(x_{t+s} \mid x_t, x_{t-1}, ...)] = \\ &MSE[\hat{E}(x_{t-s} \mid x_t, x_{t-1}, ..., y_t, y_{t-1}, ...)] \end{aligned} \qquad (41)$$

Equivalently, we say that *x is exogenous in the time series sense with respect to y* if (41) holds. A third expression meaning the same thing is that *y is not linearly informative about future x.*

Granger causality can be tested assuming a particular autoregressive lag length p and estimate

$$\begin{aligned} x_t = \; & c_1 + \alpha_1 x_{t-1} + \alpha_2 x_{t-2} + ... + \alpha_p x_{t-p} + \\ & + \beta_1 y_{t-1} + \beta_2 y_{t-2} + ... + \beta_p y_{t-p} + u_t \end{aligned} \qquad (42)$$

by OLS. We then conduct an F test of null hypothesis

$$H_0 : \beta_1 = \beta_2 = ... = \beta_p = 0 \qquad (43)$$

Consequently, if the result of the test is greater than the 5% corresponding critical value, then we reject the null hypothesis that y does not Granger-cause x.

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