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TESTING LONG-RUN PURCHASING POWER PARITY UNDER EXCHANGE RATE TARGETING

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ABSTRACT

The present paper exploits the idea that empirical estimates of the long-run PPP relationship may compound two distinct influences coming from the behavior of market participants and policy makers when the latter are targeting the exchange rate. This tends to bias tests of long-run PPP against its acceptance. The validity of the theoretical arguments is assessed by drawing on the experience of two European Union countries, Greece and France for the post-Bretton Woods period. Estimation biases due to the omission of policy effects are found to be significant only in the case of Greece. For France, our test results provide evidence bearing on the effectiveness of the competitive disinflation strategy pursued by the French authorities.

Keywords: Long-run PPP; exchange rate targeting; intervention policy; multivariate cointegration

JEL classification: C32; F31; F33

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1. Introduction

Purchasing power parity (PPP) is one of the most extensively analyzed relationships in the international finance literature (see e.g. the recent survey studies by Breuer, 1994; Froot and Rogoff, 1995; Rogoff, 1996). In its relative version it states that changes in nominal exchange rates should equal inflation differentials or, equivalently, that real exchange rates should be constant. The underlying notion is that deviations from the parity represent profitable commodity arbitrage opportunities which, if exploited, will tend to bring the exchange rate towards the parity. Although, as a description of reality PPP is clearly an oversimplification, it has been used as at least a long-run relationship in a large number of open economy models (see *inter alia* MacDonald and Taylor, 1992; Froot and Rogoff, 1995) since the return to a floating exchange rate regime in the early 1970s. Yet, its empirical verification as either a short-run or a long-run relationship has generally been rather poor (Froot and Rogoff, 1995). In particular, the failure of PPP to hold in the short run became obvious in the years immediately following the move to floating rates in March 1973, and few proponents of PPP would now argue for continuous PPP. Instead, PPP is seen as a parity condition linking relative prices and the exchange rate in the long run.

Even in its long-run form, PPP was often difficult to establish empirically. Various explanations for the failure of long-run PPP based on theoretical or statistical arguments have, therefore, been put forward. The main theoretical arguments regard the nature of shocks in the economy and problems related to transaction costs. Shocks that hit the real exchange rate and are permanent can lead to non-acceptance of mean reversion for the real exchange rate. Nevertheless, many authors question the existence of these types of shocks over the post-Bretton Woods period. Similarly, transaction costs can drive a wedge between relative prices and the exchange rate, precluding the finding of long-run PPP if these costs do not follow a stationary process.

Statistical arguments explain an apparent rather than a real failure of PPP in the long run. These arguments are mainly related to the low power of the statistical tests used and to measurement errors in prices. In the presence of such problems, rejection of long-run PPP may be considered as a statistical artifact (Michael *et al.*, 1994). For example, empirical studies, which use longer samples or panel data or more powerful testing techniques, usually result in the acceptance of PPP as a long-run condition.

In this paper we offer an alternative explanation for the apparent failure of long-run PPP by considering the effects of policy behavior. The basic idea is that coefficient estimates of the long-run relationship between the exchange rate and relative prices may compound two distinct influences, one coming from the behavior of market participants and the other from the behavior of policy makers in case the latter are targeting the exchange rate[1]. Market participants, on the one hand, tend to establish PPP in the long run, although their short-run behavior may be influenced by interventions of the monetary authorities in the foreign exchange market. The monetary authorities, on the other hand, may undertake interventions in the market to support an exchange rate rule that they may follow. If the short-run behavior of market participants is actually affected by interventions which, in turn, are governed by a policy rule, then testing for long-run PPP by examining the behavior of exchange rates and relative prices alone, would produce a long-run coefficient between these two variables which depends on the policy rule parameter. As a result, the long-run coefficient would in general be different from unity even though long-run PPP holds, i.e. the true coefficient is unity.

The validity of the theoretical arguments is assessed empirically by considering the performance of two European economies, those of Greece and France, for the recent floating exchange rate period. The choice of this sample was motivated by the fact that the monetary authorities of the two countries have been pursuing - although in a different institutional setting - an implicit or explicit exchange rate target for the whole or part of the period analyzed. The empirical results confirm the theoretical postulates and offer support for long-run PPP when policy effects are taken into account. When estimation biases regarding the long-run effects are present in simple tests of PPP, as in the case of Greece, accounting for short-run policy effects allows one to draw correct inference as regards the validity of long-run PPP. Only when the authorities follow a PPP rule, can estimation biases of the long-run effects be avoided. This turned out to be the case for France, even though the French authorities aimed *a priori* to improve competitiveness, along with targeting the nominal exchange rate. In all cases, however, interventions of the authorities in the foreign exchange market

[1] A somewhat similar argument involving the existence of a policy rule as regards the management of interest rates by monetary authorities was developed and used by McCallum (1994) to explain inefficiency of the forward market for foreign exchange as implied by rejections of the unbiasedness hypothesis.

may influence the short-run behavior of the exchange rate, whereas in the long run we expect PPP to hold.

From an econometric point of view, the theoretical analysis points to the need for very careful modeling of the short-run effects when testing for long-run PPP. If the short-run effects are basically different from the long-run effects, the explicit specification of the former is probably crucial for a successful estimation of the latter and of the time path to equilibrium (see also Juselius, 1995). This is particularly the case when the adjustment is very slow, as is the adjustment to PPP. For this reason, long-run PPP is tested as an equilibrium relationship using the Johansen multivariate cointegration technique (Johansen, 1988; Johansen and Juselius, 1990). The model specification advocated by the technique allows for different short-run and long-run effects. In addition, the Johansen technique allows for possible interactions in the determination of the variables and thus neither relative prices nor the exchange rate have to be considered *a priori* exogenous. It also allows testing for the alternative versions of PPP as presented in the literature (expressed as linear restrictions on the long-run parameters) and does not consider any of them as given.

The rest of the paper is organized as follows: Section 2 briefly surveys the recent literature on PPP. Section 3 develops the arguments and describes the theoretical model. Section 4 presents the data set and reports the empirical results. Finally, Section 5 summarizes and concludes.

2. Recent evidence on long-run PPP

Most of the recent empirical research on PPP tests for the validity of this relationship during the post-Bretton Woods period. In the 1980s, empirical studies commonly failed to support PPP, as the hypothesis of mean reversion for the real exchange rate was outperformed by the random walk hypothesis (Roll, 1979; Mark, 1990). This inability to detect mean reversion was often interpreted as indicating that real exchange rates are governed by permanent shocks (Grilli and Kaminsky, 1991; Stockman, 1990), but analyses of long historical data sets or calculations of statistical measures of shock persistence suggested that shocks to real exchange rates have a finite life (Lothian, 1997; Olekalns and Wilkins, 1998). In addition, empirical work testing whether real exchange rates are stationary processes with a permanent shift (Wu, 1997) provided positive evidence for long-run PPP, whereas in other studies the long run was interpreted as a period of sufficient length for the effects of real shocks to die out (Breuer, 1994; MacDonald, 1995). Thus, the view gradually emerged in the recent literature that PPP holds in the long run, even though it is difficult to establish it empirically. This view is based on both econometric and economic considerations.

The econometric studies in favor of this view adopt one of the following four main approaches: i) the use of long time series, ii) the use of panel data and methods, iii) the use of tests with improved power and iv) the use of advanced cointegration techniques. Froot and Rogoff (1995) emphasized the need for long time series[2], whereas Edison (1987), Abuaf and Jorion (1990) and Lothian and Taylor (1996), among others, applied univariate techniques to long samples covering one or two centuries and found evidence in favor of PPP. This approach has been criticized, however, because it combines regimes of fixed and floating rates and can be subject to large sample biases (Engel, 2000). Moreover, it does not always provide positive evidence (Cuddington and Liang, 2000).

The second approach advocates panel tests using data from a large number of countries. Most studies provide positive evidence for PPP for most of the countries analyzed (Jorion and Sweeny, 1996; Papell, 1997; Papell and Theodoridis, 1998; Bayoumi and MacDonald, 1999; and Fleissig and Strauss, 2000), although not to the same extent for each country (Koedijk *et al.*, 1998). However, in some cases the panel unit root tests did not produce strong evidence of PPP (Abuaf and Jorion, 1990; Frankel and Rose, 1996). This approach has also been criticized on the grounds that real exchange rates may be cross-sectionally dependent (O`Connell, 1998).

The third approach suggests the use of tests with improved power and advanced time series techniques. Following this approach, Abuaf and Jorion (1990), Sarno and Taylor (1998) and Kuo and Mikkola (1999) provide results that give support to long-run PPP. In the same line, Lothian and Taylor (1997) state that rejection of PPP reflects the low power of unit root tets. Cheung and Lai (1998) test for PPP by using sequential unit root tests which extend the ADF test to account for possible breaks in the real exchange rate series and they argue that permanent shocks are not relevant in PPP analysis over the current float. They also argue that the puzzling behavior of real exchange rates stems from long-memory dynamics (Cheung and Lai, 2001). When

[2] In particular, they showed that if the real exchange rate follows a stationary AR(1) process and the true half life of PPP is 3 years, it would take 72 years of data to reject the unit root hypothesis.

these dynamics are properly accounted for, strong evidence of non-monotonic mean reversion can be unveiled (Cheung and Lai, 2000). Michael *et al.* (1997) and Baum *et al.* (2001) apply exponential "smooth transition" threshold autoregressive models (ESTAR) and provide evidence of a mean-reverting nonlinear process for sizeable deviations from PPP.

A general criticism of the above methods stems from the fact that tests for real exchange rate stationarity impose *a priori* the symmetry and proportionality restrictions on exchange rates and prices and, thus, can bias PPP tests towards not accepting mean reversion. Since these restrictions need not hold empirically in the presence of measurement errors in prices (Cheung and Lai, 1993; MacDonald and Moore, 1996), the appropriate method may be to test for cointegration between exchange rates and domestic and foreign prices. Although single-equation methods – mainly the Engle and Granger method - almost always fail to find cointegration between the three variables (Taylor, 1988; Ardeni and Lubian, 1991), system methods – mainly the Johansen method - provide evidence of cointegration (Edison *et al.,* 1997; Juselius, 1995; MacDonald, 1993; Cheung and Lai, 1993). The system cointegration methods also deal with the problem of the endogeneity of the variables.

Another strand of the literature has focused on economic explanations for the persistence of deviations from PPP that cause the inability to unveil the mean-reverting tendency of real exchange rates. In an effort to provide an economic rationale, Sercu *et al.* (1995) develop a dynamic equilibrium model of real exchange rate determination and base the slow parity reversion to market frictions that impede commodity trade. If transaction costs are high enough to produce a substantial "band of inaction" within which deviations from PPP are not arbitraged away by market forces, then a linear model will fail to support convergence (Taylor and Peel, 2000). Thus, all linear models can bias testing towards finding slow convergence or random walk and the appropriate models to test for PPP are the ESTAR-type models.

To conclude, recent work indicates that a consensus seems to have been formed among researchers towards accepting relative PPP as a long-run approximation to the true arbitrage condition in the goods markets. The approach we take below can be thought of as contributing to this consensus since it offers an alternative explanation based on policy considerations for the inability to establish empirically PPP in certain cases, when in fact it is valid as a long-run relationship.

3. The theoretical framework

As already indicated, few proponents of PPP would at present argue for instantaneous PPP. PPP is rather seen as a long-run parity condition, while in the short run, we would observe deviations from it driven by the different speed of adjustment of prices and the exchange rate. According to this view, goods prices are sticky, whereas exchange rates, being financial asset prices, are known to adjust quickly to nominal shocks. A number of factors have been reported to underlie these different speeds of adjustment, most of which are related to the characteristics of the goods and asset markets.

The present paper concentrates on a different factor, that is policy behavior, and, in particular, intervention of the monetary authorities in the foreign exchange markets when the authorities follow an exchange rate rule. Intervention often aims at a certain exchange rate level that is not necessarily the PPP level and this may cause *ceteris paribus* short-run deviations from PPP. The idea that intervention in the foreign exchange market may lead to short-run divergences from PPP goes back to Cassel who indicated that the government may "bid up the price of foreign exchange above the PPP by demanding a certain amount of foreign currency irrespective of price" (see Officer, 1976). In Cassel´s writings, however, the purpose of government intervention was the procurement of foreign exchange as a substitute for capital inflows rather than as a means of influencing the course of the exchange rate, in contrast to the situations during the period under consideration.

Omission of intervention effects can be shown to have interesting implications when intervention supports a policy rule and is effective in influencing the exchange rate and creating deviations that are above or below PPP for long periods. Each time PPP tends to be reestablished by market forces, exchange rates guided by intervention move to sustain the deviation. However, as will be analyzed in detail below, the intervention policy cannot be pursued indefinitely[3]. Therefore, PPP will be established in the long run, although these policy-induced deviations may influence our ability to detect it. In particular, if the policy is sustained for a sufficiently long time and the sample is relatively small, these deviations may be mistakenly detected as a factor influencing the long-run behavior of the real exchange rate. Thus, not only could

[3] Because either the policy goal has been achieved or intervention is no longer successful in postponing the return to PPP.

omission of intervention effects distort the pattern of market adjustment to PPP but, more importantly, such omissions could introduce estimation biases to the long-run effects.

One simple way of isolating the short-run effects of intervention is to incorporate into the dynamics of the PPP model a variable, which accounts for intervention. To illustrate, consider the following simplified dynamic PPP model:

$$\Delta s_t = \lambda(s_{t-1} - \alpha - p_{t-1} + p^*_{t-1}) + \mu\Delta R_t \qquad \text{with} \qquad \lambda<0, \quad \mu>0 \tag{1}$$

where s,p and p^* are the exchange rate, defined as units of domestic currency per unit of foreign currency, the domestic and foreign price levels respectively, all expressed in logs, R is the (net) holding of foreign assets by the central bank and Δ denotes the first difference operator. Equation (1) may be viewed as a restricted version of an error correction equation for the exchange rate derived from a Vector Autoregression (VAR) that involves the exchange rate and relative prices and in which policy behavior influences the exchange rate in the short run.

The first term on the right-hand side of equation (1) represents short-run deviations from PPP which are corrected through time at a speed given by λ. In other words, the term captures the influence of market forces that tend to establish PPP in the long run. Intervention, as proxied by ΔR[4], can alter the market adjustment path towards PPP by exerting an impact on the short-run dynamics of the exchange rate, and this is captured by the second term on the right-hand side of equation (1). The coefficient μ is expected to be positive in the general case: when, for instance, the central bank sells foreign currency ($\Delta R < 0$), the exchange rate appreciates ($\Delta s < 0$) and this can be considered as the direct effect of the intervention.

In reality, there may be a secondary effect on the movement of the exchange rate coming from the reaction of market participants to the policy signal. Market participants can either strengthen the intervention effect (in case they themselves start selling foreign currency), or weaken it (in case they follow the opposite strategy). Therefore, in the extreme case that market participants follow the opposite strategy, and this secondary effect dominates the direct intervention effect, we would expect μ

[4] Changes in reserves may not correspond perfectly to interventions for a number of reasons (Neely, 2000). Their use, however, has been a common strategy given the scarcity of data on interventions.

to have a negative sign. Market participants′ decision depends on the credibility of the intervention policy, which is influenced by a number of factors, one of them being the time during which this policy has been implemented. In particular, the longer intervention policy is implemented, the less credible it is expected to become. For example, if the intervention, which targets, let's say, an appreciation of the domestic currency is continued for a long time period, the market may believe that the room for manoeuvre of the central bank has narrowed as foreign exchange reserves have fallen below a certain threshold, so the central bank is not willing to continue this policy. Thus, there will be expectations of depreciation of the domestic currency and market participants will reverse their behavior and start buying foreign currency (selling domestic currency).

Equation (1) implies that relative PPP is valid in the long run. Indeed, the long-run relationship corresponding to equation (1) is:

$$s = \alpha + p - p^{*} \tag{2}$$

Let us assume now that ΔR is determined by a policy rule consistent with the monetary authorities′ objectives. A possible rule that has been found to have merits in describing actual short-run intervention strategies (Artus, 1976, Sarno and Taylor, 2001) can be formalised as:

$$\Delta R_t = k_1(s_t - \bar{s}_t) + k_2 \Delta s_t \ , \qquad k_1,\ k_2 < 0 \tag{3}$$

with $$\bar{s}_t = \beta_0 + \beta_1(p_t - p_t^{*}) \tag{4}$$

According to equation (3), the intervention policy aims to reduce deviations of the exchange rate from a target rate $(\bar{s})$ set in terms of relative prices as shown in equation (4), and at moderating exchange rate changes. The first component in the policy reaction function (equation (3)) means that monetary authorities direct their intervention towards maintaining a particular target level of the exchange rate. Thus, if the current value of the exchange rate is higher (lower) than its target value, the central bank will intervene by selling (buying) foreign exchange, i.e., $k_1 < 0$. The target level of the exchange rate need not be the PPP level. Thus, for example, for a country that uses its exchange rate as a disinflationary means, $\beta_1 < 1$. This policy creates forces that produce one-way deviations from PPP – in this case below PPP.

The second component in equation (3) can encompass two intervention strategies (Quirk, 1977): first, countering large fluctuations of the exchange rate on a very short-term basis and second, "leaning against the wind", which implies resistance to market forces of longer duration. In the "leaning against the wind" strategy, the monetary authorities have the option of responding to market pressures by moderating exchange rate movements to varying extents: they can allow market pressures to be absorbed by large exchange rate fluctuations and limited use of reserves for intervention or vice versa.

What insights can be gained from the system of equations (1), (3) and (4) as to the validity of long-run PPP? Substitution of equations (3) and (4) into equation (1) yields

$$\Delta s_t = \lambda(s_{t-1} - \alpha\ \beta_0 - p_{t-1} + p^*_{t-1}) + \mu k_1(s_t - \beta_0 - \beta_1\ (p_t + p^*_t)) + \mu k_2 \Delta s_t \tag{5}$$

which describes the short-run dynamics of the exchange rate. The static solution of equation (5) is:

$$s = \gamma_0 + \gamma_1(p - p^*) \tag{6}$$

where $\gamma_0 = (\lambda\alpha + \mu\beta_0) / (\lambda + \mu k_1)$ and $\gamma_1 = (\lambda + \mu k_1\beta_1) / (\lambda + \mu k_1)$.

Equation (6) is the static equivalent of equation (5), although it is not a long-run relationship since it incorporates short-run policy effects. However, this is the equation that we test as a long-run relationship if we do not account for the intervention effects. It can be easily seen that the coefficient γ_1 is a function of the policy rule parameter β_1. As demonstrated by our analysis, the estimated γ_1 coefficient would take the value 1 only if $\beta_1 = 1$, i.e. the authorities follow a PPP rule. In the general case (when $\beta_1 \neq 1$), we would be unable to accept the hypothesis of long-run PPP, even though it is true. Thus, the empirical finding that $\gamma_1 \neq 1$ merely reflects that $\beta_1 \neq 1$, i.e. that there is a policy rule different from PPP, and that this policy rule affects market behavior for a sufficiently long time, so that it is mistakenly interpreted as a long-run influence.

This finding has interesting implications. Testing for PPP by deriving equation (6) as the long-run relationship without accounting for short-run policy effects would in general result in an estimated γ_1 coefficient different from unity and this could be interpreted, at first sight, as evidence against long-run PPP. However, such an interpretation would be false. If we isolate the short-run policy effects – by using equation (1) – then we will be able to detect long-run PPP empirically when this is valid and/or obtain more sensible estimates of the speed of mean reversion.

4. Empirical evidence

In this section we test for long-run PPP by drawing on the experience of two European Union countries, Greece and France, during the post-Bretton Woods period[5]. Both countries were members of the European Monetary System (EMS); France also participated in the Exchange Rate Mechanism (ERM) of the EMS since its inception, while Greece became a member in March 1998. Despite this institutional dissimilarity, a common element in these countries' macroeconomic strategies is that both targeted the exchange rate for a part or the whole of the period analyzed.

In Greece, a tight exchange rate policy was pursued from the late 1980s, restricting the depreciation of the drachma against other European currencies to less than would be needed to accommodate inflation differentials (Detragiache and Hamann, 1999; see also Figure 1). In terms of the policy rule (equation 4) of the previous section, one would *a priori* expect the coefficient β_1 to be less than one. The above policy is thought to have been successful in imposing discipline on Greek firms regarding the containment of their costs and their pricing behavior. By the end of 1998, inflation as measured by the consumer price index, had declined to 3.9% from 19% at the beginning of the 1990s. The role of exchange rate targeting was central in the disinflation process, while other policies – incomes and fiscal – were less restrictive, at least in the early part of the period. Over the same period, wage developments were characterised by swings in the rate of change of real unit labor costs, which, at least initially, may have had a negative influence on the formation of inflationary

[5] Our theory was also tested against the experience of two Nordic countries – Sweden and Finland – as they both followed a "strong currency" exchange rate policy in the 1980s. In both cases, the results were similar to those obtained for Greece and are not reported here, but are available on request.

expectations leading to a higher degree of inflation inertia. Additionally, the fiscal deficit was on average above 10% of GDP and only after 1996 did it record a rapid reduction, helped by the fall in inflation and interest rates.

France achieved disinflation successfully under the discipline of the exchange rate commitment in the ERM. After a period of expansionary policies and sharp devaluations of the franc, France undertook disinflation in 1983, a year considered a turning point in its macroeconomic policy. The logic of French disinflation, that came to be known as “competitive disinflation” (Trichet, 1992), was that, by achieving and preserving lower inflation than its trading partners, France would experience improved competitiveness and would finally return to conditions of full employment, after having absorbed the initial cost of disinflation. In terms of equation (4) above, this policy would be translated into a parameter β_1 greater than one. Disinflation was achieved by 1987 and inflation remained at low levels thereafter (cf. Figure 2). Both incomes and fiscal policies were behind the success of disinflation. Incomes policy contributed to speeding up the reduction of inflation through disindexation, i.e. through linking nominal wages to target inflation that was lower than actual inflation.

Fiscal policy was also under control throughout the disinflation period and the fiscal (general government) deficit moved from 3.2% of GDP in 1983 to 1.2% in 1989 – although it increased thereafter and returned again to a downward path after 1994. Disinflation was associated with modest competitiveness gains within the EU which, however, were offset by the movements of the dollar and the large devaluations of the lira and the pound after the EMS crisis. Overall, disinflation in France has been successful in eliminating the French inflation differential, but has not been accompanied by improved competitiveness as intended (Blanchard and Muet, 1993).

In the empirical analysis that follows, we use effective exchange rates for the drachma and the franc (geometrically weighted averages of bilateral exchange rates, where the weights are shares in the external trade of each country accounted for by the ten major trading partners) and weighted foreign price levels, the weights being the same as those for the exchange rate. This allows us to deal with the bilateral bias problem - the large correlation that may exist between bilateral exchange rate movements and may lead to imprecise estimates. The price variables are measured by the consumer price index (CPI), given that CPIs are broadly similar as far as coverage is concerned (OECD, 1994).[6] Quarterly seasonally unadjusted data for the period 1972(1) to 1997(4) are used. The data come from *International Financial Statistics* and the *Direction of Trade Statistics* of the International Monetary Fund (see Appendix A). Effective estimation periods are reduced so as to accommodate the lag structure of the estimated models. All variables are expressed in logs.

The empirical work is performed in two steps. In the first step, testing for long-run PPP is done in the traditional (in terms of model specification) framework of a third-order VAR, which allows for interdependence of prices and the exchange rate, by using the Johansen cointegration technique. We estimate this simple model in the three variables, in order to verify empirically that omission of the intervention effects yields an equation between the exchange rate and prices in the form of equation (6). In the second step, we assess the validity of the theoretical hypotheses advanced in Section 3. We therefore test for long-run PPP (as a cointegrating relationship) in a model, which takes into account the short-run policy effects. If PPP (equation (2)) is established empirically with this modification, we can conclude that the rejection of PPP in the first step was in fact due to the estimation biases stemming from the omission of short-run intervention effects.

The empirical work emphasizes that the lack of empirical support for PPP may partly be due to lack of an appropriate specification of the short-run dynamics of the PPP model. In particular, it emphasizes the importance of allowing for different long-run and short-run effects, the latter including intervention effects, so that the error (or equilibrium[7]) correction term of the model ensures that in equilibrium, i.e. in an economy with no changes, interventions or shocks, the model is consistent with the parity.

[6] The use of "harmonised" CPIs would be more appropriate but unfortunately these are available only for European Union countries since 1995.

[7] Clements and Hendry (1995) advocate that what are known as "error correction models" should be called "equilibrium correction models", based on the observation that in such reparameterizations the long-run information terms known as "error corrections" first introduced by Davidson *et al.* (1978) may play the opposite role when the equilibrium changes. However, the traditional terminology is adopted in the present paper.

4.1 The econometric methodology

Equation (2) defines long-run equilibrium in the goods market, whereas equations (1) and (3) to (5) describe the short-run dynamics of the variables of interest, in a very simplified world. When used as a basis for empirical modeling they have to be modified, so that the stochastic properties of the data are taken into account. In addition, there might be other factors not specified by the theory that are relevant to understanding the variation in the series, i.e. policy changes, exogenous shocks, or structural breaks. In order to take into account such problems presented in applied work, the analysis follows the "General to Specific" methodology[8], in which the time-series properties of the data play a dominant role. In a system context, it advocates as an initial step, the construction of a congruent unrestricted vector autoregression (VAR), which can be considered as an adequate representation of the joint distribution of the observed series, the so-called data generating process (DGP) (see, *inter alia,* Hendry and Mizon, 1993).

In the VAR framework, the number of the cointegrating relationships between the variables can be defined following the Johansen procedure (Johansen, 1988). The procedure suggests a reparameterization of the initial VAR, in the familiar vector error correction (VEC) form:

$$\Delta x_t = \sum_{t=1}^{p-1} \Pi_i \Delta x_{t-i} + \Pi x_{t-p} + \psi D_t + v_t \tag{7}$$

where x_t is an N×1 vector of the time series of interest v_t~IN(0,Σ) and D_t contains a set of conditioning variables. The order of the VAR, p, is assumed finite and the parameters Π_i, Π and ψ are assumed constant. Π is the matrix of the long-run responses and, if there exist r cointegrating relationships between the variables, is of reduced rank r < N. In this case, Π can be expressed as the product of two N×r matrices α and β' : Π = α β' where β contains the r cointegrating vectors and α is the

[8] For a detailed analysis of the methodology, see, *inter alia*, Hendry (1995) and Spanos (1986).

loadings matrix, which contains the coefficients with which the cointegrating relationships enter the equations modeling Δx_t .

Johansen and Juselius (1990) and Johansen (1995) provide the test statistics to define the rank r of the matrix Π and show that testing for linear restrictions on either the parameters of the cointegrating vectors or their loadings is allowed given that the matrices α and β' are not unique. Therefore, specific meaningful economic restrictions concerning the elements of the matrices α and β can be tested and not imposed *a priori*. In the present case, certain linear restrictions on the elements of the matrix β can test theoretical hypotheses for the long-run behavior of the variables, in particular the hypotheses of symmetry and proportionality. Certain restrictions on the elements of the matrix α may imply weak exogeneity of the variables *p, p*, s* with respect to the long-run parameters. In particular, zero restrictions on the elements of the matrix α test whether or not the cointegrating vectors enter the equations of the system[9] (i.e. whether or not the variables are error-correcting).

4.2 Long-run PPP in Greece

Testing for long-run PPP is first undertaken for Greece. We estimate a three-dimensional VAR of the form of equation (7) for the vector $x = (s, p^*, p)$, where the variables are as defined in Section 3. Following the "General to Specific" methodology, the model is estimated using five lags for the variables initially, with a constant and seasonals included in the conditioning variables set D_t. However, likelihood ratio tests indicated the number of lags to be 4 in the final model. As shown in Figure 1, the data exhibit large fluctuations, which can be partly considered as resulting from policy changes. Such fluctuations often lead to empirical rejection of the statistical assumptions as well as loss of efficiency. In fact, while residual correlation and heteroscedasticity could not be rejected by the VAR specification, the normality of the residuals was, possibly due to non-constant parameters as indicated by the plots of the relevant Chow tests. These features supported the inclusion of impulse dummies to account for any structural breaks observed in the sample

[9] For a presentation of the concept of weak exogeneity see *inter alia* Ericsson (1994); for testing for weak exogeneity in the cointegration framework, see Johansen (1995).

period[10]. Three impulse dummies were included, one for the oil price shock and two for the devaluations of the drachma in January 1983 and October 1985. They all turned out to be significant in the system, whereas their absence would mean non-normal residuals.

The statistical properties of the residuals of the final VAR specification are reported in Table B1 in Appendix B. The diagnostics do not indicate any serious mis-specification and therefore the VAR can be considered as a congruent statistical representation of the data[11]. The VAR satisfies the statistical assumptions required for the Johansen technique and thus we can go on with the cointegration analysis. The outcomes of the maximum eigenvalue and trace statistics are reported in Table 1. According to both likelihood ratio tests, there is evidence of one cointegrating vector. In addition, the plot of the recursively estimated maximum eigenvalue indicates a cointegrating relationship with constant parameters.

The estimated coefficients of the cointegrating vector (shown in Table 1) indicate that it does not necessarily express a PPP relationship. Further testing is, therefore, required. Table 1 also presents the outcomes of the likelihood ratio test statistics for alternative hypotheses concerning the specification of the estimated cointegrating vector. The specific hypotheses tested are those of symmetry and proportionality, which imply "weak" and "strong" PPP, respectively. According to the test outcomes, strong PPP is not accepted by the data set, whereas weak PPP is accepted for a relationship of the form $s - 0.68\left(p - p^{*}\right)$.

In the lower part of Table 1 the results of weak exogeneity tests of the variables with respect to the weak PPP parameters are also reported.[12] The results indicate that, consistent with the small country assumption, foreign prices are weakly exogenous and that adjustment to PPP operates mainly through prices. However, the relevant

[10] Inclusion of dummies is preferable to an enlargement of the system, as advocated by Clements and Mizon (1991).

[11] The tests on the VAR residuals indicate an autocorrelation problem, which cannot be solved by adding more lags of the variables. Moreover, inclusion of more lags does not change the test results on weak and strong PPP. Given that inclusion of more lags would reduce the available degrees of freedom and that Gonzalo (1994) has demonstrated the robustness of the Johansen procedure, we make no further modeling changes, but we should keep in mind that our results should be qualified in this respect.

[12] In the present case (where there is evidence for one cointegrating vector), these tests are essentially tests for the significance of the cointegrating vector when used as error correction term in the VECM reparameterization.

hypotheses are accepted/rejected at the margin, given that the values of the test statistics are very close to each other and very close to the critical values (at the 5 % level of significance); thus, they cannot be considered as conclusive for the weak exogeneity status of the variables.

4.3 Long-run PPP under a policy rule

In Section 3, it was demonstrated that, if the authorities direct their intervention policy towards targeting a specific exchange rate level, then the parameter of the long-run PPP relationship can be expected to be a function of the policy rule parameter. In Greece, as already indicated, the authorities used the appreciation of the real exchange rate as a disciplinary means in order to achieve disinflation, and this has possibly influenced the short-run dynamics of the exchange rate. Consequently, the results of the tests presented above could be due to the fact that we do not specify adequately the short-run dynamics of the model.

We test this hypothesis by introducing policy effects in the model; to this end, we include into the system an intervention variable, proxied by the change in foreign exchange reserves ΔR_t. As indicated in equation (1), we assume that these interventions have influenced the short-run behavior, whereas in the long run we expect the levels to adjust according to the steady state PPP. We therefore include current and lagged values[13] of ΔR_t in the conditioning variables set D_t[14]. By doing so, we estimate a conditional system in which ΔR_t is assumed to be weakly exogenous with respect to the cointegrating parameters (estimation of the full system would imply ΔR_t to be included in the vector x_t).

The estimation of the VAR is made using 4 lags for x_t, as above, and D_t also includes a constant and seasonal and event specific dummies. The short-run effects of ΔR_t turned out to be significant for the system (F-statistic for the null hypothesis that ΔR_t is not included in the VAR specification: $F(3,77)=3,47$). The inclusion of ΔR_t did not alter the stochastic properties of the VAR system as indicated by the results of the diagnostic tests reported in Table B2 in Appendix B.

[13] Following the "General to Specific" methodology, we included initially 4 lags of ΔR_t, but then kept those, which were significant (i.e. ΔR_t).
[14] We follow a similar specification to that of Juselius (1995).

The results from the Johansen cointegration analysis are presented in Table 2. There is again evidence of one cointegrating vector[15] with the coefficients having the theoretically expected signs and magnitudes. In addition, the graph of the recursive estimate of the maximum eigenvalue of the system indicates parameter constancy of the cointegrating relationship.

Next, the weak exogeneity status of ΔR_t is assessed. This is done because, on the basis of the findings of Johansen (1992), maximum likelihood cointegration analysis in a conditional model provides an identical estimator to that based on a full system, if the conditioning variables (in our case ΔR_t) are weakly exogenous for the cointegrating parameters (see also Banerjee, Hendry and Mizon, 1996; Urbain, 1995, *inter alia*)[16]. The weak exogeneity test is essentially a test for the significance of the cointegrating vector when used as an error correction term in a single equation, modelling the behavior of ΔR_t. ΔR_t turns out to be weakly exogenous for the cointegrating vector parameters (t-statistic for the error correction term: 0.775).

Table 2 also presents the outcomes of the tests of the theoretical hypotheses which concern the cointegrating parameters. According to them, the symmetry hypothesis cannot be rejected for a vector of the form $s - 0.98\left(p - p^*\right)$, implying that it possibly expresses a PPP relationship since the relevant coefficient is close to unity. Indeed, the proportionality hypothesis cannot be rejected in the system, contrary to the results of the previous system. This finding supports the existence of estimation biases, whose removal by the inclusion of ΔR_t allows us to disentangle long-run PPP[17].

The results of the weak exogeneity tests lead us to comfortably accept the hypothesis of weak exogeneity of foreign prices, while this hypothesis is rejected for the exchange rate and domestic prices. The size and sign of the $\alpha_{1k,\ (k=1,\ldots 3)}$ coefficients in Table 2 - α_{11}, α_{12} and α_{13} are the coefficients with which the first and unique

[15] As advocated by the magnitude of the estimated eigenvalues and the values of the two likelihood ratio test statistics.

[16] As Banerjee *et al.* (1996) state: "if the feedback (from policy rules) is not related to deviations from long-run equilibria, policy variables can be weakly exogenous, and then asymptotically efficient inference on the parameters of the agents' model can be obtained without the need to analyze the policy makers' decision rules simultaneously".

[17] With the particular form chosen for the policy rule (equations 3 and 4), it is clear by looking at equation (6) why omission of the intervention variable from the estimation prevents us from finding empirical support for the strong form of PPP (provided it holds true). Had we specified the policy rule differently (e.g. in terms of changes in *s* and *p-p**), a possible non-acceptance of the strong form of PPP would still reflect estimation bias.

cointegrating vector enters the equations modelling Δp^*, Δp and Δs, respectively - indicate that domestic prices change in response to disequilibrium in a way as to bring the system back to equilibrium. By contrast, the exchange rate does not seem to move towards eliminating any deviations from equilibrium but rather moves in a way that does not accommodate inflation differentials so as to promote disinflation. As noted above, this policy was successful in bringing down inflation in Greece, an argument corroborated by our finding that prices took the burden of adjustment to PPP. Moreover, exchange rate policy appears to have acquired credibility, as indicated by the positive sign of the intervention variable in the VAR equation for the exchange rate.

4.4 PPP tests for France

Long-run PPP was tested for France following similar steps. Testing was initially attempted within the framework of a fourth-order three dimensional VAR. The VAR also includes two impulse dummies to account for the effects of the first oil price shock and the realignment of the central parity of the franc in the ERM in June 1982 (which was the largest in the period under consideration). The VAR turns out to be well-specified as indicated by the diagnostics reported in Table B1 in Appendix B. The cointegration results presented in Table 1 support the existence of one cointegrating vector. The outcomes of the tests for the alternative hypotheses indicate that weak PPP is accepted for a specification of the form s -1.22(p-p*) that is, with an estimated β_1 coefficient numerically greater than one. As discussed above, we would *a priori* expect such a result, had the policy of competitive disinflation followed by the French authorities been successful. However, testing for strong PPP suggests that this hypothesis is also accepted by the data.

The extended system (which includes intervention effects) does not provide notably different results regarding the acceptance of the strong version of PPP. The intervention variable, included finally with no lags, turns out to be significant for the exchange rate equation (t-statistic: -2.29) although it appears with a negative sign. This suggests that exchange market interventions of the French authorities did not convince market participants about the feasibility of exchange rate targeting and, at the same time, of controlling inflation in the context of their competitive disinflation strategy which aimed to create a permanent deviation from PPP. Instead, as confirmed by our test results, long-run PPP was established *ex post*. Finally, weak exogeneity tests indicate that adjustment to equilibrium comes through exchange rate movements (related to the importance of non-ERM currencies in the effective exchange rate of the franc) and not through prices whose movements were influenced by the anti-inflationary fiscal and incomes policies pursued.

5. Concluding remarks

PPP has long attracted the interest of economists and has served as a useful building block in a large number of open economy macroeconomic models. Yet, even in its long-run version, PPP has often proved difficult to establish empirically. Several explanations based on theoretical and statistical arguments have been put forward to explain empirical failures of PPP. In this paper we offer an alternative hypothesis by considering the relevance of long-run PPP in a framework that allows for influences caused by the implementation of an exchange rate rule by the authorities when they are targeting the exchange rate. The novelty of our approach is in emphasizing that the coefficient estimates of long-run PPP may compound two distinct effects coming from the behavior of policymakers intervening in the foreign exchange market in support of a policy rule and of market participants engaging in goods arbitrage. Thus, there is a potential bias towards not accepting PPP even as a long-run relationship.

In the empirical part of the paper we have illustrated the interaction between policy behavior and market adjustment by estimating the PPP model with data from two EU countries, Greece and France, for the post-Bretton Woods period. The analysis, which employed a multivariate cointegration technique, involved testing for hypotheses maintained in previous studies, namely symmetry and proportionality, as well as for the weak exogeneity status of the variables. The empirical results support the validity of our theoretical arguments. In particular, the results are very supportive of long-run PPP in the case of Greece, once policy effects are taken into account. Indeed, accounting for foreign exchange market intervention in the model led to the acceptance of strong PPP in the long run, accompanied by meaningful exogeneity properties for the variables. For France, the results revealed that biases due to policy effects are not as important as in the case of Greece, confirming that the competitive disinflation policy pursued by the French authorities succeeded only in maintaining competitiveness in the long run rather than improving it.

Our results can be viewed as complementary to the growing body of recent empirical evidence of long-run PPP over the recent float. The new element introduced by our analysis is the investigation of the short-run PPP dynamics and the mechanism establishing this arbitrage condition in the long run in the presence of an intervention policy under exchange rate targeting, which potentially biases empirical tests of long-run PPP. However, our findings show that policy behavior, while affecting short-run adjustment to PPP and our ability to uncover long-run PPP, cannot prevent the long-run tendency towards purchasing power parity.

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Table 1

Cointegration analysis of the VARs with no intervention effects.

Greece			France		
Testing for the $\mathbf{\Pi}$ rank.					
Eigenvalues	Max. Eigen.	Trace	Eigenvalues	Max. Eigen.	Trace
0.2085	23.16*	36.79**	0.2535	29.25**	42.11**
0.1275	13.51	13.63	0.1187	12.64	12.86
0.0012	0.12	0.12	0.0022	0.19	0.19
s	p*	p	s	p*	p
Standardized $\boldsymbol{\beta}'$ cointegrating vector					
1.000	1.333	-0.974	1.000	1.215	-1.222
Standardized $\boldsymbol{\alpha}$ coefficients					
0.001	-0.012	0.020	-0.316	-0.018	-0.039

	Greece	France
Testing for structural restrictions		
H_0	χ^2 statistic (p-value)	
Weak PPP	χ^2 (1) = 3.87 [0.049]*	χ^2 (1) = 0.09 [0.756]
Strong PPP	χ^2 (2) = 10.79 [0.004]**	χ^2 (2) = 2.95 [0.229]
Weak exogeneity testing		
	χ^2 statistic (p-value)	
s	χ^2 (2) =6.12 [0.049] *	χ^2 (3) = 21.20 [0.000]**
p*	χ^2 (2) =5.87 [0.053]	χ^2 (3) = 6.49 [0.091]
P	χ^2 (2) =6.38 [0.041] *	χ^2 (3) = 9.97 [0.019]*

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

Table 2

Cointegration analysis of VARs including policy effects

Greece			France		
Testing for the **Π** rank.					
Eigenvalues	Max. Eigen.	Trace	Eigenvalues	Max. Eigen.	Trace
0.1943	21.60*	39.04**	0.2577	29.80**	42.61**
0.1552	14.85	17.44*	0.1191	12.69	12.81
0.0057	0.57	0.57	0.0012	0.12	0.12
s	p*	P	s	p*	P
Standardized **β'** cointegrating vector					
1.000	1.265	-0.980	1.000	1.201	-1.210
Standardized **α** coefficients					
0.012	-0.013	0.029	-0.333	-0.006	-0.033

	Greece	France
Testing for structural restrictions		
H_0	χ^2 statistic (p-value)	
Weak PPP	$\chi^2(1) = 3.76$ [0.510]	$\chi^2(1) = 0.07$ [0.795]
Strong PPP	$\chi^2(2) = 4.30$ [0.116]	$\chi^2(2) = 2.48$ [0.290]
Weak exogeneity testing		
	$\chi2$ statistic (p-value)	
S	$\chi^2(3)= 7.88$ [0.049] *	$\chi^2(3)= 21.78$ [0.000] **
p*	$\chi^2(3)= 5.62$ [0.132]	$\chi^2(3)= 6.25$ [0.101]
P	$\chi^2(3)=20.98$[0.000]**	$\chi^2(3)= 9.87$ [0.019]*

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

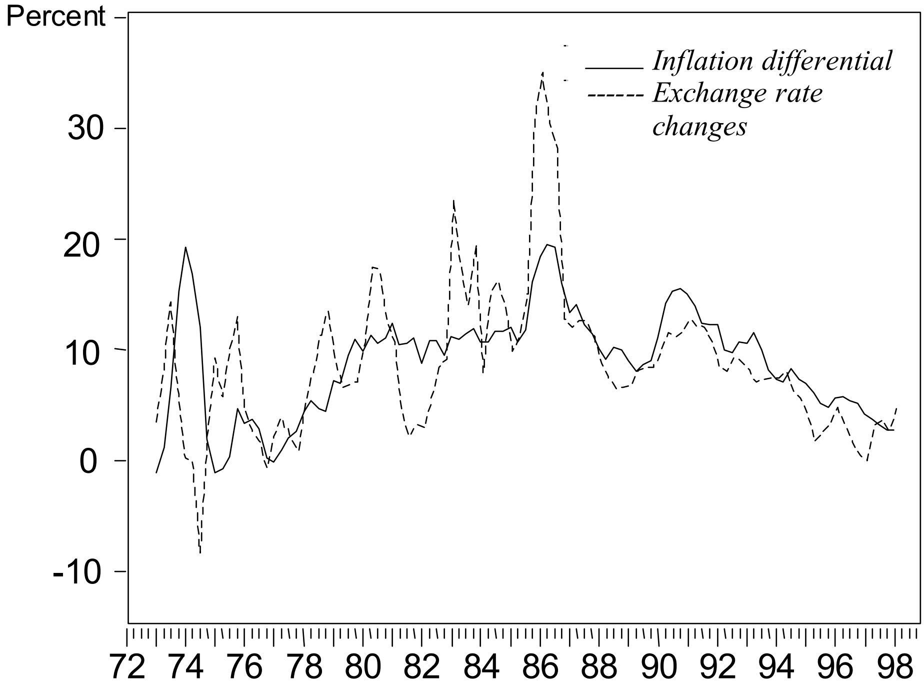

Figure 1. Inflation differential and exchange rate changes: Greece

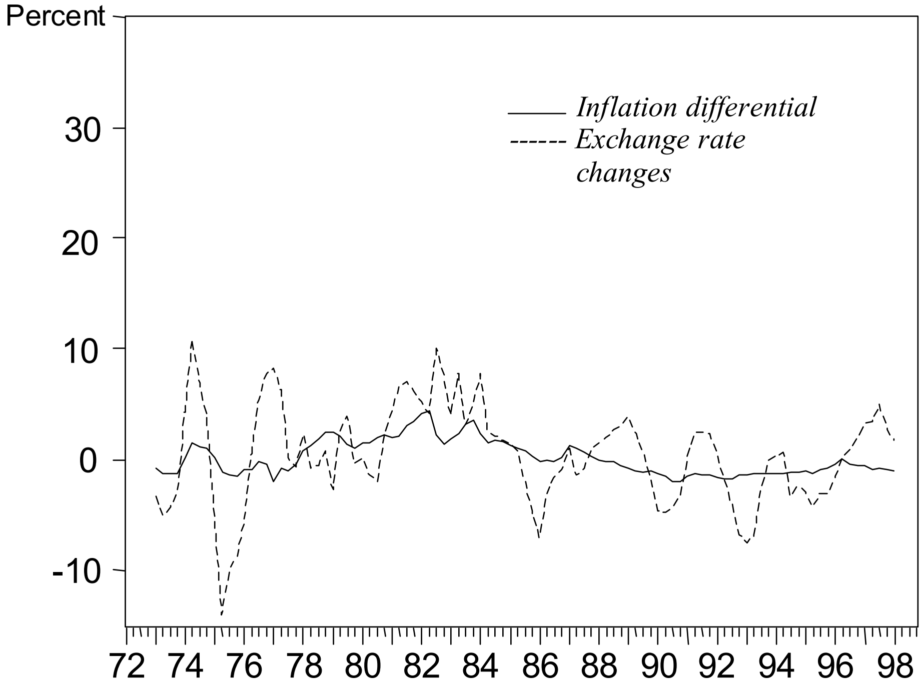

Figure 2. Inflation differential and exchange rate changes: France

Appendix A.

Data definitions and sources

- P: Domestic prices. Consumer price index, 1980 = 100 (IMF, International Financial Statistics (IFS), line 64).
- S: Effective exchange rate. Weighted average of bilateral exchange rates against the ten main trading partners of each country (IFS, line number rf), the weights being the shares in the external trade (imports plus exports), for the years 1994-1997 (IMF, Direction of Trade Statistics).

  For Greece, the weights are: Germany: 0.25, Italy: 0.25, France: 0.11, UK: 0.09, Netherlands: 0.08, USA: 0.06, Spain: 0.05, Belgium: 0.05, Japan: 0.04, Austria: 0.02.

  For France, the weights are: Germany: 0.24, Italy: 0.14, UK: 0.13, Belgium: 0.12, Spain: 0.10, USA: 0.10, Netherlands: 0.07, Sweden: 0.04, Japan: 0.04, Austria: 0.02.
- P*: Foreign prices. Weighted average of consumer price indices (IFS, line 64) of the ten main trading partners, the weights being the same as for the effective exchange rate.
- R: Foreign exchange reserves (IFS, line 11 minus line 16c).

Appendix B.

Table B1

Misspecification tests for initial VARs

	Greece	France
	Equation residuals	
Autocorrelation		
s	AR 1- 5 $F(5, 76) = 2.45$ [0.041] *	AR 1- 5 $F(5, 77) = 1.783$[0.126]
p*	AR 1- 5 $F(5, 76) = 0.95$ [0.456]	AR 1- 5 $F(5, 77) = 1.593$ [0.172]
p	AR 1- 5 $F(5, 76) = 3.24$ [0.010] *	AR 1- 5 $F(5, 77) = 1.691$ [0.147]
Normality		
s	$\chi^2(2) = 6.15$ [0.046] *	$\chi^2(2) = 5.30$ [0.070]
p*	$\chi^2(2) = 4.00$ [0.135]	$\chi^2(2) = 2.46$ [0.292]
p	$\chi^2(2) = 0.46$ [0.792]	$\chi^2(2) = 1.33$ [0.515]
Conditional heteroscedasticity		
s	ARCH 4 $F(4, 73) = 1.21$ [0.314]	ARCH 4 $F(4, 74) = 0.42$ [0.791]
p*	ARCH 4 $F(4, 73) = 2.41$ [0.056]	ARCH 4 $F(4, 74) = 1.03$ [0.396]
p	ARCH 4 $F(4, 73) = 4.15$ [0.005]**	ARCH 4 $F(4, 74) = 0.51$ [0.727]
	VAR residuals	
Autocorrelation	AR1-5 $F(45,190) = 1.77$ [0.004]**	AR1-5 $F(45,193) = 1.53$ [0.026] *
Normality	$\chi^2(6) = 7.42$ [0.283]	$\chi^2(6) = 7.73$ [0.259]
Cond. hetero/city	$F(144,305) = 1.21$ [0.084]	$F(144,311) = 1.22$ [0.081]

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

Table B2

Misspecification tests for the VARs including policy effects

		Greece	*France*
		Equation residuals	
	Autocorrelation		
s		AR 1- 5 F(5, 73) = 2.38[0.046]*	AR1-5 F(5,76) = 1.17 [0.331]
p*		AR 1- 5 F(5, 73) =0.61 [0.689]	AR1-5 F(5, 76) = 1.49 [0.202]
P		AR 1- 5 F(5, 73) = 2.72 [0.026]*	AR1-5 F(5, 76) = 1.76 [0.132]
	Normality		
s		χ^2 (2) = 5.68 [0.059]	χ^2 (2) = 7.26 [0.026]*
p*		χ^2 (2) =2.83 [0.243]	χ^2 (2) = 2.51 [0.285]
p		χ^2 (2) =0.79 [0.676]	χ^2 (2) = 1.75 [0.418]
	Conditional heteroscedasticity		
s		ARCH 4 F(4, 70) =1.05 [0.387]	ARCH 4 F(4,73) = 0.26 [0.906]
p*		ARCH 4 F(4, 70) =1.96 [0.110]	ARCH 4 F(4,73) = 0.99 [0.418]
p		ARCH 4 F(4, 70) = 2.82 [0.031]*	ARCH 4 F(4,73) = 0.46 [0.762]
		VAR residuals	
Autocorrelation		AR 1-5 F(45,181) =1.43 [0.055]	AR 1-5 F(45,190) = 1.33[0.096]
Normality		χ^2 (6) = 5.80 [0.446]	χ^2 (6) = 10.05 [0.122]
Cond. Hetero/city		F(144,288) =1.09 [0.278]	F(144,305) = 1.13 [0.189]

Note: * and ** indicate rejection of the null hypothesis at the 5% and 1% level of significance, respectively.

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