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PERMANENT AND TRANSITORY DYNAMICS IN HOUSE PRICES AND CONSUMPTION: CROSS-COUNTRY EVIDENCE

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Permanent and Transitory Dynamics in House Prices and Consumption: Cross-Country Evidence *

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Abstract

In this paper a small-scale macroeconomic system is estimated in the framework of a common trends model, in order to explore the dynamic interactions between real house prices, consumption expenditure and output in the US and major European economies. The results point to important differences across countries, with long-run house price effects on consumption only for France, Germany and the US. However, some interactions between house prices and consumption are detected in all countries at shorter horizons. Evidence of international comovements in the common trend component of house price dynamics is also found.

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1 Introduction

The aim of this paper is to characterize the dynamic interactions among house prices and consumption, separating permanent movements from transitory fluctuations. Though dating back at least to Modigliani (1971), the interest in empirical estimation of wealth effects on consumption expenditure has been recently revived following several episodes of boom and bust cycle in both the stock market and the housing markets. The importance of disentangling permanent ("trend") from transitory ("cyclical") changes in wealth is pointed out by Lettau and Ludvigson (2004) who empirically identify permanent and transitory elements in US household net worth, and investigate how they are related to consumer spending. Their main finding is that the bulk of fluctuations in household wealth are dominated by the transitory component, and therefore are unrelated to aggregate consumer spending, since the latter reacts only to permanent wealth movements.

Starting from the Lettau-Ludvigson insight, we use the econometric framework of the common trends model of King, Plosser, Stock and Watson (1991) to investigate the impact of permanent and temporary real house price dynamics on consumption using quarterly data over the period 1979-2007 for the major European economies (France, Germany, Italy, Spain and the UK).[1] In order to benchmark the findings, the analysis has also been carried out for the US. For each country, permanent and transitory movements in house prices and consumption are estimated within a three-variable system including, beside real house prices and private final consumption expenditure, also output. In the common trends framework, the permanent component of the endogenous variables bears the meaningful economic interpretation of long-run forecast, conditional on the information contained in the system. Moreover, by means of a minimal set of identifying assumptions, we are able to give economic content to the (two) permanent innovations driving the system and study their individual dynamic effect on house prices and consumption at different horizons. In fact, the inclusion of output in the system allows for a simple way of separating supply-side (e.g. productivity) shocks, affecting output in the long-run, from output-neutral disturbances related to the demand side.

Our findings point to significant long-run effects of house price fluctuations on consumer spending only for some countries (France, Germany and the US), whereas in the countries where no long-run effect is found (Italy, Spain and the UK) house prices and consumption appear to be related in the

[1] Since most of the quarterly fluctuations in housing wealth is attributable to house price movements, we can focus on the latter variable to capture housing wealth fluctuations in our sample.

short- to medium-term horizon. Finally, a principal component analysis carried out on the permanent component of real house price dynamics provides strong evidence of global comovements across countries.

The paper is organized as follows. In the next section the econometric methodology is outlined and the data set is described. Section 3 presents the empirical results, concering the long-term (cointegration) relationships, the estimation of permanent and transitory components in the time series behaviour of the variables, the analysis of their dynamic responses to structural permanent disturbances, the characterization of the transitory fluctuations in house prices, and the investigation of the existence of common global dynamics among the house price trends of the six countries under study. Section 4 summarizes the main conclusions.

2 Econometric methodology and data

We study the interactions among house prices, output and consumption by means of three-variate country-specific models, aiming at capturing the main features of the joint dynamics of the macroeconomic variables of interest and providing an accurate identification of shocks with a different degree of persistence. To this aim, we apply the common trends methodology of King, Plosser, Stock and Watson (1991) and Mellander, Vredin and Warne (1992), exploting the long-run (cointegration) properties of the data to disentangle the permanent and transitory components in the time-series behaviour of house prices, consumption and output. In this context, the permanent component of each series bears the interpretation of a long-run forecast conditional on the information contained in the three-variable system analyzed and consistent with the long-run cointegration properties of the data.

The rest of this section outlines the econometric methodology in some detail and presents some descriptive evidence of the data used.

2.1 The common trends model

Consider a vector $\mathbf{x}_t$ of n $I(1)$ variables of interest (in the application below $\mathbf{x}_t$ includes the logs of private final consumption expenditure, real house prices, and GDP). If there exist $0 < r < n$ cointegrating relations among the variables, the following cointegrated VAR representation for $\mathbf{x}_t$ holds (deterministic terms are omitted throughout for ease of exposition)

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

where $\mathbf{\Pi}(L) = \mathbf{\Pi}_1 + \mathbf{\Pi}_2 L + ... + \mathbf{\Pi}_p L^{p-1}$ is a polynomial in the lag operator L, the $n \times r$ matrix $\boldsymbol{\beta}$ contains the cointegrating vectors (capturing long-run equilibrium relations), such that $\boldsymbol{\beta}'\mathbf{x}_t$ are stationary linear combinations of the variables, $\boldsymbol{\alpha}$ is the $n \times r$ matrix of loadings (capturing the adjustment of each variable in $\mathbf{x}$ to deviations from long-run equilibrium), and $\boldsymbol{\varepsilon}_t$ is a vector of serially uncorrelated reduced form disturbances. As shown in Mellander, Vredin and Warne (1992), the cointegrated VAR in (1) can be inverted to yield the following stationary Wold representation for $\Delta\mathbf{x}_t$:

$$\Delta\mathbf{x}_t = \mathbf{C}(L)\boldsymbol{\varepsilon}_t \tag{2}$$

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + ...$ with $\sum_{j=0}^{\infty} j \mid \mathbf{C}_j \mid < \infty$. Starting from (2) it is possible to derive the *stochastic trends* representation of $\mathbf{x}_t$, decomposing the series into a permanent (non-stationary) and a transitory (stationary) components, whereby extending the Beveridge and Nelson (1981) univariate decomposition to a multivariate framework. By recursive substitution, we obtain the following expression for the levels of the variables:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{C}(1)\sum_{j=0}^{t-1} \boldsymbol{\varepsilon}_{t-j} + \mathbf{C}^*(L)\boldsymbol{\varepsilon}_t \tag{3}$$

where $\mathbf{x}_0$ is the vector of the initial values of the series, $\mathbf{C}^*(L) = \sum_{j=0}^{\infty} \mathbf{C}_j^* L^j$ with $\mathbf{C}_j^* = -\sum_{i=j+1}^{\infty} \mathbf{C}_i$, and $\mathbf{C}(1)$ captures the long-run effect of the reduced form disturbances in $\boldsymbol{\varepsilon}$ on the variables in $\mathbf{x}$.

The existence of cointegrating relations linking the elements of $\mathbf{x}$ imposes restrictions on the $\mathbf{C}(1)$ matrix, constraining the long-run responses of the n endogenous variables. With r cointegrating vectors, the non-stationary component of $\mathbf{x}$ can be expressed in terms of a reduced number $k = n - r$ of *common stochastic trends* as follows:

$$\mathbf{x}_t = \mathbf{x}_0 + \underset{(n\times k)}{\mathbf{A}} \underset{(k\times 1)}{\boldsymbol{\tau}_t} + \mathbf{C}^*(L)\,\boldsymbol{\varepsilon}_t \tag{4}$$
$$\text{with} \quad \boldsymbol{\tau}_t = \boldsymbol{\tau}_{t-1} + \boldsymbol{\psi}_t$$

where $\boldsymbol{\tau}_t$ is a k-element vector random walk and $\boldsymbol{\psi}_t$ contains the k innovations to the stochastic trends, i.e. the *permanent shocks.* Matrix $\mathbf{A}$ captures the impact of the (common) stochastic trends on each variable in $\mathbf{x}$. The common trends representation in (4) not only separates the permanent component of $\mathbf{x}$ from the transitory component but also attributes the permanent component to a limited number (k) of permanent disturbances that can possibly be separately identified and whose individual dynamic effects on $\mathbf{x}$ can be studied by means of impulse response analysis and forecast error variance decompositions.

Permanent component of the series. The analysis of the properties of the common trends model in (4) begins by noting that the *permanent component*, $\mathbf{x}_t^P$, can be easily obtained from the long-run forecast for $\mathbf{x}$ since in the long-run only the stochastic trends have an influence on the levels of the non-stationary endogenous variables (the transitory component $\mathbf{C}^*(L)\,\boldsymbol{\varepsilon}_t$ being stationary). Hence:

$$\mathbf{x}_t^P = \lim_{i\to\infty} E_t\,\mathbf{x}_{t+i} = \mathbf{x}_0 + \mathbf{A}\,\boldsymbol{\tau}_t \tag{5}$$

capturing the values to which the series are expected to converge once the effect of the transitory shocks have died out. Thus, no assumption on the correlation between permanent and transitory innovations and on the structural economic nature of the shocks are needed to estimate the permanent component of the series. However, if we are interested also in estimating the long-run effect of *each individual* structural permanent disturbance in $\boldsymbol{\psi}$ and the *dynamic response* of each variable in $\mathbf{x}$ to such shocks, then complete identification of the nk elements of $\mathbf{A}$ is necessary. In the presence of multiple common trends ($k > 1$), the decomposition of the stochastic permanent component $(\mathbf{A}\,\boldsymbol{\tau}_t)$ into a matrix of loadings $\mathbf{A}$ and a vector of common stochastic trends $\boldsymbol{\tau}_t$ cannot be based on purely statistical grounds but requires some economic assumptions.

Identification of structural permanent disturbances. To carry out this step, and obtain an economically meaningful interpretation of the dynamics of the variables of interest from the reduced form representations in (2) and (3), the vector of reduced form disturbances $\boldsymbol{\varepsilon}$ must be transformed into a vector of underlying structural shocks, some of which with *permanent* effects on the level of $\mathbf{x}$ and some with only *transitory* effects. Let us denote this vector of i.i.d. structural disturbances as $\boldsymbol{\varphi}_t \equiv \begin{pmatrix} \boldsymbol{\psi}_t \\ \boldsymbol{\nu}_t \end{pmatrix}$, where $\boldsymbol{\psi}$ and $\boldsymbol{\nu}$ are subvectors of k and r elements respectively, with $k = n - r$. The structural form for the first difference of $\mathbf{x}_t$ is:

$$\Delta\mathbf{x}_t = \boldsymbol{\Gamma}(L)\boldsymbol{\varphi}_t \tag{6}$$

where $\boldsymbol{\Gamma}(L) = \boldsymbol{\Gamma}_0 + \boldsymbol{\Gamma}_1 L + \ldots$. Since the first element of $\mathbf{C}(L)$ in (2) is $\mathbf{I}$, equating the first term of the right-hand sides of (2) and (6) yields the following relationship between the reduced form and the structural shocks:

$$\boldsymbol{\varepsilon}_t = \boldsymbol{\Gamma}_0\boldsymbol{\varphi}_t \tag{7}$$

where $\mathbf{\Gamma}_0$ is an invertible matrix. Hence, comparison of (6) and (2) shows that

$$\mathbf{C}(L)\mathbf{\Gamma}_0 = \mathbf{\Gamma}(L)$$

implying that $\mathbf{C}(1)\mathbf{\Gamma}_0 = \mathbf{\Gamma}(1)$. In order to identify the elements of $\boldsymbol{\psi}_t$ as the permanent shocks and the elements of $\boldsymbol{\nu}_t$ as the transitory disturbances the following restriction on the long-run matrix $\mathbf{\Gamma}(1)$ must be imposed:

$$\mathbf{\Gamma}(1) = (\mathbf{A} \quad \mathbf{0}) \tag{8}$$

with $\mathbf{A}$ being the $n \times k$ matrix in (4). The disturbances in $\boldsymbol{\psi}_t$ are then allowed to have long-run effects on (at least some of) the variables in $\mathbf{x}_t$, whereas the shocks in $\boldsymbol{\nu}_t$ are restricted to have only transitory effects.

From (6), the common trends representation of $\mathbf{x}_t$ in structural form is derived as

$$\begin{aligned}
\mathbf{x}_t &= \mathbf{x}_0 + \mathbf{\Gamma}(1)\sum_{j=0}^{t-1}\boldsymbol{\varphi}_{t-j} + \mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t \\
&= \mathbf{x}_0 + \mathbf{A}\sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j} + \mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t \\
&= \mathbf{x}_0 + \mathbf{A}\,\boldsymbol{\tau}_t + \mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t
\end{aligned} \tag{9}$$

where the first equality makes use of (8), $\boldsymbol{\tau}_t$ is the k-variate random walk defined in (4), and $\mathbf{\Gamma}^*(L)$ is defined analogously to $\mathbf{C}^*(L)$ in (3).

As shown in detail by Stock and Watson (1988), King, Plosser, Stock and Watson (1991) and Warne (1993), the identification of separate permanent shocks requires a sufficient number of restrictions on the long-run impact matrix $\mathbf{A}$ in (9). Part of these restrictions (rk) are provided by the r cointegrating vectors, requiring that

$$\boldsymbol{\beta}'\mathbf{A} = \mathbf{0} \tag{10}$$

A second set of $k(k+1)/2$ restrictions on the elements of $\mathbf{A}$ is obtained by equating the two representations of $\mathbf{x}$ in (3) and (9), yielding

$$\mathbf{C}(1)\,\boldsymbol{\varepsilon}_t = \mathbf{A}\,\boldsymbol{\psi}_t \tag{11}$$

(a restatement of the fact that the long-run impact of $\boldsymbol{\varepsilon}$ is only due to the permanent structural innovations $\boldsymbol{\psi}$). From this relation it follows that (imposing $E(\boldsymbol{\psi}_t\boldsymbol{\psi}_t') = \mathbf{I}$)

$$\mathbf{C}(1)\,\mathbf{\Sigma}\,\mathbf{C}(1)' = \mathbf{A}\,\mathbf{A}' \tag{12}$$

where $\mathbf{\Sigma}$ is the variance/covariance matrix of the VAR innovations $\boldsymbol{\varepsilon}$. The remaining $k(k-1)/2$ restrictions needed for (exact) identification of $\mathbf{A}$ have then to be derived from economic theory and can take the form of zero restrictions on some of its elements (e.g. in the case of long-run neutrality assumptions). Once identification od $\mathbf{A}$ is achieved, estimates of the structural permannet disturbances may be obtained from (11) as

$$\boldsymbol{\psi}_t = (\mathbf{A}'\mathbf{A})^{-1}\mathbf{A}'\mathbf{C}(1)\,\boldsymbol{\varepsilon}_t$$

and, from the moving average representation in (6), impulse responses and forecast error variance decompositions may be calculated to gauge the relative importance of each permanent innovation in determining fluctuations of the endogenous variables.[2]

Transitory dynamics. An important property of the permanent-transitory decomposition obtained from the common trends model is that the transitory component $\mathbf{x}_t^{TR}$ is determined by both permament ($\boldsymbol{\psi}_t$) and transitory shocks ($\boldsymbol{\nu}_t$). From (9):

$$\mathbf{x}_t^{TR} = \mathbf{\Gamma}^*(L)\,\boldsymbol{\varphi}_t = \mathbf{\Gamma}_\psi^*(L)\,\boldsymbol{\psi}_t + \mathbf{\Gamma}_\nu^*(L)\,\boldsymbol{\nu}_t \tag{13}$$

where the first component $\mathbf{\Gamma}_\psi^*(L)\,\boldsymbol{\psi}_t$ gives the contribution of permanent innovations to the overall transitory fluctuations (dynamics "along the attractor"), while the vector $\mathbf{\Gamma}_2^*(L)\boldsymbol{\upsilon}_t$ measures the contribution of the transitory disturbances, linked to the process of adjustment towards long-run equilibrium (dynamics "towards the attractor"). The two components have a fundamentally different economic interpretation. The adjustment dynamics have the error correction process as generator, and therefore are "disequilibrium" fluctuations. On the contrary, the dynamics along the attractor may be related to the overshooting of the variables to permanent innovations, capturing the transitional dynamics which take place after a shock to the common trends of the system; since along the attractor the cointegration relationships are satisfied, the dynamics along the attractor are "equilibrium" fluctuations. In our application, following Proietti (1997) and Cassola and Morana (2002), we disentangle the two components of the transitory dynamics in real house prices, to provide some additional insights into the nature of house price fluctuations.

[2] In addition, nr restrictions are needed to identify r separate transitory shocks in $\boldsymbol{\nu}_t$. To this aim, a set of $kr + r(r+1)/2$ restrictions are provided by the orthogonality conditions $E\left[\boldsymbol{\psi}_t\boldsymbol{\nu}_t'\right] = \mathbf{0}$ and $E\left[\boldsymbol{\nu}_t\boldsymbol{\nu}_t'\right] = \mathbf{I}_r$. Then additional $r(r-1)/2$ restrictions, grounded in economic theory, are to be imposed for exact identification. In the current application only one transitory disturbance is present and no additional restrictions are needed for identification.

2.2 Data and descriptive statistics

For each country, we specify a three-variable system including (all in logs) private final consumption expenditure (c), an index of real house prices (h) and GDP (y) sampled at a quarterly frequency. The countries under investigation are: France, Germany, Italy, Spain, the UK, and the US. The source of the data is $OECD$; in particular, house price data are extensively described and analyzed in Girouard, Kennedy, van den Noord and André (2006). The sample period ranges from 1978(1) to 2007(4) with the only exceptions of Spain (starting from 1980(1)) and Italy (ending in 2007(3)). From the important perspective of the prevailing system of housing finance (that may importantly affect the relation between fluctuations in house prices and consumption expenditure), our sample includes the three lowest-ranked countries (France, Germany and Italy) in the IMF (2008) "Mortgage Market Index", the highest (the USA) and two countries in intermediate position (Spain and the UK).[3]

Figure 1 displays the yearly rates of growth of the three variables for all countries, and Table 1 offers some descriptive statistics on the same variables. The behaviour of real house prices shows remarkable differences across countries. In particular, wide fluctuations occurred in Italy, Spain and the UK (with standard deviations between 8% and 9%), whereas the other countries feature less pronounced fluctuations (5.4% for France, 3.3% for the USA, and 2.7% for Germany[4]). Also the contemporaneous correlations between house price growth and consumption and GDP growth display different patterns: high (positive) correlation in Spain and the UK (around 0.7 with consumption growth and 0.6 with GDP growth), no correlation in Italy, and intermediate results in the remaining countries. Such evidence points to possibly important cross-country differences in the dynamics linking house prices to consumption expenditure and output. In order to fully characterize the long- and short-run dynamic interactions among the variables, disentangling permanent from transitory components, we now turn to the application of the common trends model outlined above.

[3] On a scale ranging from 0 to 1, the IMF index is 0.98 for the US, 0.58 and 0.4 for the UK and Spain respectively, and around 0.25 for the remaining countries.

[4] In the case of Germany, the results must be taken with caution since they may be affected by the unification occurred in 1990. In the econometric analysis below we allowed for shifts in the variables after 1990; moreover, results on the shorter post-unification sample (1991-2007) are qualitatively similar to those obtained on the full sample.

3 Empirical results

For each country the initial specification of the three-variable VAR system in levels has been set to five lags and then progressively reduced, testing each step by means of a battery of specification tests. The final specifications of the unrestricted reduced form model in levels feature two lags for Spain, three for Germany and Italy, four for the UK and the US; only in the case of France five lags have been retained in the model.

3.1 Cointegration analysis

To test for the existence of long-term relationships among the investigated variables, different criteria have been jointly employed. Johansen's (1988) trace test has been used to assess the number of valid cointegrating relationships, and the Johansen reduced rank regression approach has been implemented to estimate the cointegrating vectors in the cointegrated VAR (1). We also relied on the Granger representation theorem, whereby the presence of error correcting behavior within a set of nonstationary $I(1)$ variables is a sufficient condition for cointegration, while the presence of cointegration within a set of variables necessarily implies the existence of an error-correction mechanism. Therefore, we looked also at the statistical significance of the elements of $\boldsymbol{\alpha}$ in (1) as additional evidence of cointegration. Finally, standard information criteria have been used to further evaluate the estimated cointegrated vector error-correction model against the unrestricted alternative.

The results of cointegration analysis are reported in Table 2. Overall, evidence of one cointegrating vector can be obtained for all countries, albeit clear-cut results can be attained only by considering the error-correcting properties of the variables and the information criteria computed with and without imposing cointegration rank and identification restrictions. Tests based on the trace statistic (as shown by the p-values reported in the upper part of the table) clearly support cointegration for France, Italy, the US, and, to a lesser extent, for Spain, whereas for Germany and the UK the evidence in favour of cointegration comes from the strongly significant estimates of the error-correcting coefficients. Moreover, in all cases both the AIC and BIC information criteria point to the cointegrated model as the preferred one.

Unrestricted estimates of the cointegrating vector ($\boldsymbol{\beta}$) and the error-correction coefficients ($\boldsymbol{\alpha}$) are shown in the middle part of Table 2, whereas in the lower part of the table appropriate restrictions (in all cases supported by the reported likelihood ratio test) are imposed on the structure of $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ (the restricted cointegrating vectors are shown in Figure 2). Two groups of countries emerge from the results. On the one hand, Italy, Spain and the UK are characterized by a long-run relationship which involves only consumption and output with no role for house prices.[5] This finding points to the lack of long-term effects of real house price movements on consumption expenditure, though the possibility of sizeable short- to medium-term wealth effects on consumption stemming from house price dynamics is still allowed, as shown by the impulse response analysis carried out below. In general, a permanent change in house prices will have a positive wealth effect on landlords and owner-occupiers, and a negative income effect on tenants and on prospective first-time buyers, so that an aggregate non-zero effect on consumption is detectable if a change in house prices entails a redistribution between agents with different marginal propensities to spend. In addition, since housing wealth may be collateralizable and therefore used to finance consumption, an increase in housing wealth could relax borrowing constraints and boost consumption.[6] Under the latter respect, while the lack of long-term effects of house prices on consumption is not very surprising for Italy and Spain, for the UK the evidence is somewhat at odds with the larger role of collateralized consumer credit in this country. However, recent evidence by Benito, Thompson, Waldron and Wood (2006) points to a gradually weakening association between UK house prices and consumer expenditure in recent years that is consistent with our findings.

On the other hand, in France, Germany and the US, also the house price variable enters the cointegration relationship together with consumption and output, pointing to long-term housing wealth effects. The magnitude of the estimated long-run elasticities is larger in Germany (0.26) and smaller (but strongly statistically significant) in France and the US (0.13 and 0.16 respectively). Finally, the estimated error-correction coefficents in $\boldsymbol{\alpha}$ show that house prices strongly react to deviations from the equilibrium relations in France and the US, suggesting that house price dynamics contains a quantitatively important transitory component that dies out in the long-run (Lettau and Ludvigson, 2004).

3.2 Permanent and transitory components

The existence of one cointegrating relationship among three $I(1)$ non stationary variables implies the presence of two distinct sources of shocks having

[5] In the case of Italy and Spain, given the (1, 0, -1) structure of the cointegrating vector, there is evidence of stationarity of the consumption-output ratio, whereas in the case of the UK the long-run relationship between c and y is not homogeneous.

[6] Buiter (2008) provides a clear theoretical analysis of the potential wealth effects on consumption (or lack thereof) due to house price changes.

permanent effects on at least some of the variables, and one transitory shock.[7] In terms of the common stochastic trend representation in (4), the permanent component of the series is driven by a bivariate random walk process of the form:

$$\begin{pmatrix} \tau_t^1 \\ \tau_t^2 \end{pmatrix} = \begin{pmatrix} \mu^1 \\ \mu^2 \end{pmatrix} + \begin{pmatrix} \tau_{t-1}^1 \\ \tau_{t-1}^2 \end{pmatrix} + \begin{pmatrix} \psi_t^1 \\ \psi_t^2 \end{pmatrix} \tag{14}$$

where $\boldsymbol{\mu}$ is a vector of constant drift terms, added to the model in estimation, and the levels of the variables are decomposed into a permanent and a transitory component as follows:

$$\begin{pmatrix} c_t \\ h_t \\ y_t \end{pmatrix} = \begin{pmatrix} c_0 \\ h_0 \\ y_0 \end{pmatrix} + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \\ a_{31} & a_{32} \end{pmatrix} \begin{pmatrix} \tau_t^1 \\ \tau_t^2 \end{pmatrix} + \mathbf{\Gamma}^*(L) \begin{pmatrix} \psi_t^1 \\ \psi_t^2 \\ \nu_t \end{pmatrix} \tag{15}$$

where ν_t is a purely transitory disturbance, and the elements of matrix $\mathbf{A}$, a_{ij}, capture the long-run effect of the two permanent disturbances (ψ_t^1 and ψ_t^2) on the endogenous variables. As mentioned in the preceding section, disentangling the permanent from the transitory components of consumption, house prices and output does not call for any restriction on the elements of $\mathbf{A}$, since the former is the long-run conditional forecast, determined by the joint effect of the two permanent shocks. However, in order to estimate the long-run effect of each individual disturbance and study the dynamic response of c, h and y to ψ_t^1 and ψ_t^2, we have to achieve the complete identification of the six elements a_{ij}.

To this aim, the first two sets of restrictions corresponding to (10) and (12) above, derived from the (restriced form of the) cointegrating vector and from the fact that the long-run impact of the reduced form innovations is entirely due to the permanent disturbances, provide five restrictions, leaving only one additional identifying assumption to be imposed. To achieve complete identification of $\mathbf{A}$ we impose a long-term output neutrality restriction, whereby one of the permanent disturbances, ψ_t^2, is assumed not to have a long-term effect on output y: this amounts to imposing $a_{32} = 0$ in (15). When the cointegrating vector includes only consumption and output, as for Italy, Spain and the UK, the neutrality restriction holds for consumption as well (i.e., also $a_{12} = 0$) This identifying assumption is consistent with the interpretation of the first permanent shock (ψ_t^1) as mainly a supply-side disturbance related to the engines of long-term economic growth (determining the long-run behaviour of output, consumption, and possibly real house prices), whereas ψ_t^2 has only short- to medium-term effects on output but can permanently affect real house prices and possibly consumption expenditure.

[7] In our system $n = 3$, $r = 1$, and the number of common stochastic trends $k = 2$.

The latter disturbance can capture shocks to the housing user cost, particularly to its interest rate sensitive component, i.e. the mortgage interest cost.[8] Changes in short term rates determined by changes in the monetary policy stance will impact on mortgage repayment costs and therfore on housing demand and prices. Through the aggregate demand channel, interest rate disturbances may be expected to have output-neutral effects in the long-run. Under this assumption, estimation of the common trends model in (15) is carried out, yelding the long-run effects of permanent shocks (i.e. the elements a_{ij}) reported in Table 3. Finally, for each variable, the permanent component can be constructed as, in the case of house prices:

$$\hat{h}_t^P = h_0 + \hat{a}_{21}\hat{\tau}_t^1 + \hat{a}_{22}\hat{\tau}_t^2$$

capturing the long-run effects on h of the two identified permanent disturbances, and bearing the interpretation of the (conditional) forecast of house prices over a long-term (infinite) horizon, when all transitory fluctuations in house prices have vanished. The transitory component is then simply computed as $\hat{h}_t^{TR} = h_t - \hat{h}_t^P$.

The results in Table 3 show that the permanent component of house prices, h_t^P, is determined almost entirely by the output-neutral permanent disturbance ψ_2 in France, Germany and the US (the estimates of the a_{21} element being not significant), whereas only a weak effect of ψ^1 can be detected for Italy. In Spain and the UK, instead, both permanent shocks have a strong long-run impact on house prices. A further common feature of France, Germany and the US (consistent with the presence of h in the cointegrating vector) is the significant long-run effect of ψ^2 (which basically drives house prices) on consumption. Table 4 and Figure 3 display the essential features of the estimated transitory (or, in a commonly used but not entirely appropriate terminology, "cyclical") components of consumption, house prices and output. House prices appear to be very strongly (contemporaneously) correlated with both consumption and output in Spain, the US, and France (with correlation coefficients greater than 0.8). In the UK and Germany a strong positive correlation is detected only with respect to consumption, whereas in the case of Italy both correlations are only around 0.3.

3.3 Dynamic responses to structural disturbances

To gauge the relative importance of the two identified structural permanent shocks and the transitory disturbance, a forecast error variance decomposi-

[8] Other key components of housing user costs are maintenance costs, property taxes, and expected net capital gains. See Hilbers, Hoffmaister, Banerji and Shi (2008) for additional details.

tion exercise has been carried out at different horizons, including the business cycle range (one to five years), with results reported in Table 5.

In all European countries, the permanent shock driving entirely output in the long-run (ψ^1) accounts for a large fraction of output fluctuations also over short-and medium-term horizons (the remaining fraction being accounted for mainly by the transitory shock ν), whereas in the US, at the one-quarter horizon as much as 71% of output fluctuations are attributable to ψ^2 (the permanent driving force of house prices), that still accounts for 40% of output movements at the one-year horizon. The output-neutral shock ψ^2 accounts for the bulk of house price fluctuations over long- and medium-term (business cycle) horizons in France, Germany, Italy and the US, whereas in Spain and in the UK a large fraction of long-run house prices fluctuations is explained by the permanent driving force of output (58% in Spain and as much as 78% in the UK).

As far as consumption fluctuations are concerned, in the countries where no long-run effect of house prices on expenditure is detected, the output-neutral permanent disturbance ψ^2 accounts for a non negligible fraction at a long horizon (39% in France, 26% in Germany, and 16% in the US). Figure 4, showing the impulse response function of consumption expenditure to a unitary shock in ψ^2 (with one-standard deviation error bands), confirms this finding. In particular, the effects of this shock build up gradually over time for France, Germany and the US, being already statistically significant after five quarters for Germany and the US, while for France the impact is significant already at the outset. Differently, since the impact of the shock is only transitory, the estimated dynamic response for Italy, Spain and the UK builds up for about five quarters for Italy and the UK and then fades away within three years, while a longer building up time is found for Spain (ten quarters).

Overall, two main conclusions can be drawn from the above results. First, though two mainly separate driving forces determine the long-term evolution of output and house prices, some non negligible interactions can be detected at business cycle frequencies, with the output driving force having a strong impact on house prices in some countries, especially Spain and the UK. Only for the US a relevant role of house price fluctuations for the determination of business cycle output fluctuations is found. Second, consumption and house prices do seem to be related at various frequencies, ranging from the very short-term to the long-term. In general, when countries show a long-term impact of house prices on consumption (France, Germany and the US), the latter linkage is already evident in the medium-term. Differently, when a long-term impact is absent (Italy, Spain and the UK), evidence of a linkage between consumption and house prices can however be found in the short-term.

In order to better measure the response of consumption to changes in house prices determined by the output-neutral shock, already shown in Figure 4, the impulse response function of c to ψ^2 has been normalized relatively to the impulse response of real house prices to the same disturbance. This scaling allows to measure the percentage change in consumption associated with a a 1% increase in real house prices due to the output-neutral shock. In Table 6 the normalized figures are reported for various horizons of interest, ranging from one to five year (for completeness, also the infinite-horizon eleasticities, corresponding to the coefficient on h in the estimated cointegrating vector of Table 2, are shown in the last column). At the 1-year horizon, a sizeable response of consumption to house price changes is detected for France, Germany, the US and also the UK, with the consumption elasticity falling in the range 0.12 to 0.18. Differently, a much smaller response of consumption for Italy (0.06) and Spain (0.01) can be noted. Yet, from the two-year horizon onwards, while the response of consumption is still sizeable for France, Germany, and the US, a noticeable reduction can be noted for Italy and the UK. For the latter countries and Spain, the consumption elasticity to house prices become negligible at longer horizons. Although Italy, Spain and the UK share the same feature of a zero long-run response of consumption to permanent real house price changes, the greater availability of financial products (e.g. reverse mortgages), and the stronger propensity to finance consumption through loans in the UK compared with the other two countries, could explain why consumption in the UK is more sensitive to house price changes in the short-term, than in Italy and Spain, with an elasticity over the one-year horizon very close to that of US consumption expenditure.

3.4 Transitory dynamics in house prices

As shown in the methodological section, the permanent-transitory decomposition obtained from the common trends model yields transitory components which are determined by both permanent and transitory shocks. The contribution of the two types of disturbances can be disentangled, allowing for further economic insights into the nature of transitory fluctuations in the endogenous variables. Focusing on house prices, from (15) the transitory component can be written as

$$h_t^{TR} = \Gamma_{21}^*(L)\,\psi_t^1 + \Gamma_{22}^*(L)\,\psi_t^2 + \Gamma_{23}^*(L)\,\nu_t \tag{16}$$

where $\Gamma_{2j}^*(L)$ $(j = 1, 2, 3)$ are the elements of the second row of matrix $\mathbf{\Gamma}^*(L)$ in (15). The first two terms on the right-hand side of (16) are driven by the permanent shocks and capture transitory house price fluctuations *along* the equilibrium relationship (e.g. overshooting effects in the house price equilibrium dynamics). The last term captures the contribution of the purely transitory disturbance ν_t to "cyclical" movements in house prices, capturing transitional dynamics *towards* the equilibrium relationship determined by the error-correcting properties of the data. The assessment of the relative contribution of the two components may be particularly relevant in order to establish the nature of the potential misalignment in current house prices with respect to their permanent value.

Figure 5 presents, for each country over the most recent decade (1996-2007), the overall transitory series of house prices h_t^{TR} (already shown for the whole sample in Figure 3), and the two components described above, named "transitory *equilibrium* dynamics" (capturing temporary fluctuations along the equilibrium), and "*adjustment* dynamics" (measuring the correction towards equilibrium). In all countries there is evidence of a positive misalignment of house prices with respect to their trend values in the final part of the period (the overall transitory component being positive on average), though the starting date for this process differs across countries (around 2001 in Italy and the UK, 2003 in Germany, and 2005 in France, Spain and the US). However, the most recent observations (2007) show that sizeable misalignments are still present only in France and the US. This above evidence is broadly consistent with other results in the literature pointing to a cumulated overvaluation in housing prices of about 30% since 2004, not only for the US, but also for the OECD area (Girouard, Kennedy, van den Noord and Andre 2006, Finicelli, 2007; Gros, 2007).[9]

Looking at the relative role of the transitory equilibrium and the adjustment dynamics, a sharp difference can be observed between cyclical fluctuations in house prices in the US and in the European countries for the most recent period. While for the latter countries the origin of the recent house price misalignment is related to equilibrium fluctuations, i.e. to overshooting effects along the equilibrium path, for the US current fluctuations are essentially disequilibrium dynamics induced by the error-correcting behaviour of the system. It should be noted that even for the US the findings are not consistent with a bubble interpretation of the house price misalignment. The presence of an automatic adjustment mechanism, as the one described by

[9] Yet, it should be acknowledged that the overall evidence is mixed, as, for instance, Jacobsen and Naug (2005) do not find any evidence of housing price overvaluation in the US, compared with fundamental values determined by interest rates, households income, unemployment and housing supply. Similar findings are provided by Himmelberg, Mayer and Sinai (2005) and McCarthy and Peach (2004), who also control for demographic factors.

the error correction properties, is in fact in contrast with the non stationary growth in house prices, which would be implied by a process of self-fulfilling expectations, spreading, for instance, according to social epidemics (Shiller, 2007).

3.5 Global dynamics in house prices

A growing literature on international comovement in macroeconomic variables has recently focused on the similarity of the rising house price pattern detected since the late 1990s in the G-7 countries,[10] pointing to global dynamics not only in real activity, inflation, interest rates and stock prices, but also in real house prices. Case, Goetzmann and Rouwenhorst (1999), for instance, find significant linkages between real estate prices and both local and global GDP components, suggesting that international housing price comovements are at least partially explained by common exposure to global business cycles. Similarly, Ahearne, Ammer, Doyle, Kole and Martin (2005) and Otrok and Terrones (2005) point to global real interest rate dynamics as a factor behind the international comovement in house prices. Finally, Beltratti and Morana (2008) detect comovements in G-7 house prices related to both financial and macroeconomic factors, with a key role for productivity shocks; evidence of important regional linkages is also found, especially in the euro area.

Differently from previous studies, which have investigated house price comovements directly on the actual variables, in this paper, a separate analysis for the trend ($\hat{h}_t^P$) and cyclical components ($\hat{h}_t^{TR} = h_t - \hat{h}_t^P$) has been carried out. Moreover, noting that the house price trend can be expressed as

$$\begin{aligned} \hat{h}_t^P &= h_0 + \hat{a}_{21}\hat{\tau}_t^1 + \hat{a}_{22}\hat{\tau}_t^2 \\ &\equiv h_0 + \hat{h}_t^{\tau_1} + \hat{h}_t^{\tau_2} \end{aligned}$$

the analysis has been carried out also on the two composing common trends $\hat{h}_t^{\tau_1}$ and $\hat{h}_t^{\tau_2}$. Relying on recent results of Bai (2003, 2004) and Bai and Ng (2004),[11] principal components analysis has been employed and the results reported in Table 7.

[10] Since 1999 house prices have increased at an yearly average real rate of about 5% in the US, the euro area and Canada, and to an even larger rate in the UK (close to 9%). The housing market outlook has started turning negative since early 2007, as real prices have started decreasing in the US.

[11] In particular, Bai (2003) has considered the generalization of PCA to the case in which the series are weakly dependent processes, establishing consistency and asymptotic normality when both the unobserved factors and idiosyncratic components show limited serial correlation, also allowing for heteroschedasticity in both the time and cross section dimension in the idiosyncratic components. In Bai (2004) consistency and asymptotic

There is strong evidence of international comovements in house prices, particularly in their trend dynamics. In fact, over 80% of the variance of the overall house price permanent component ($\hat{h}^P$) is explained by the first principal component, accounting for the bulk of variance for each of the house price trend series in individual countries as well, with the only partial exception of Italy (50%). Concerning the components of the house price trend due to the two permanent structural disturbances, the first principal component explains 99% of the total variance of $\hat{h}^{\tau_1}$, the common trend related to the supply-side shock (which has a long-term impact on output), also explaining a similar proportion of the variance for each of the series investigated. Hence, the evidence points to a very strong supply-side comovement for the investigated countries. Weaker comovements can be detected among the common trends related to the output-neutral shock ($\hat{h}^{\tau_2}$). For this latter set of series the first two principal components account for about 70% and 26% of total variance, respectively. The first factor captures commonalities across European countries, explaining over 90% of total variance for each country apart from France (35%), whereas the second factor is mostly related to US dynamics (88%). Given the differences in the working of mortgage markets between the US and Europe, this finding is broadly consistent with the interpretation of the output-neutral disturbance as related to changes in housing user costs.

Finally, also the transitory dynamics show some commonalities across countries, though to a smaller extent. In fact, only the first principal component points to global dynamics, accounting for about 40% of total variance with a sizeable impact on the individual series (40% to 70%), apart from Germany (19%) and Italy (4%), and the other principal components pointing to mainly regional fluctuations.

Overall, the evidence is consistent with previous results in the literature, though a more accurate understanding of comovements is achieved in the current framework. International comovement in house prices is both a trend and cycle phenomenon, but markedly stronger at the trend level. Consistent with Beltratti and Morana (2008), supply-side shocks are a key component in the explanation of this observed feature, while output-neutral shocks contribute to the explanation of both the global and regional dimensions of

normality has been derived for the case of I(1) unobserved factors and I(0) idiosyncratic components, also in the presence of heteroschedasticity in both the time and cross section dimension in the idiosyncratic components. Moreover, Bai and Ng (2004) have established consistency also for the case of I(1) idiosyncratic components. As pointed out by Bai and Ng (2004), consistent estimation should also be achieved by PCA in the intermediate case represented by long memory processes, and Monte Carlo results reported in Morana (2007) support this conclusion.

house price dynamics. From an economic point of view the results point to the global nature of technological advances, as well as to the relevance of regional valuation factors for the housing market.

4 Conclusions

The empirical analysis of the short- and long-run interactions among real house prices, consumption expenditure and output, carried out in the framework of a common trends model, detects important cross-country differences in the dynamic links between house prices and economic activity, in accord with recent evidence (IMF 2008). Most importantly, in the long-run there is no evidence for an effect of real house price fluctuations on consumption in Italy, Spain and the UK, whereas long-term wealth effects from housing are found in Germany, France and the US (with the strongest effect in the former country).

The common trends framework allows to separate permanent from transitory movements in the macroeconomic variables analyzed and study the dynamic effects of structural permanent disturbances. Overall, two main conclusions can be drawn from the dynamic analysis. First, in France, Germany, the US, and, to a lesser extent, Italy, two mainly separate driving forces determine the long-term evolution of output and house prices; however, some non negligible interactions are detected at business cycle frequencies in other countries, with the permanent (supply-side) driving force of output accounting for around 30% of house price variability in Spain and 60% in the UK at three- to five-year horizons. Second, consumption and house prices do seem to be related at various frequencies, ranging from the very short-term to the long-term. In general, when a long-term impact of house prices on consumption is found (as in France, Germany and the US), this linkage is already present in the medium-term. Differently, when a long-term impact is absent (as in Italy, Spain and the UK), there is evidence of a linkage between house prices and consumption over shorter (around one to two years) horizons. Finally, international comovements in housing prices are detected. The evidence is stronger at the trend level, and can be attributed mainly to the common trend component associated with supply-side (productivity) developments.

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Table 1
Descriptive statistics

		France				**Germany**		
		$\Delta_4 c$	$\Delta_4 h$	$\Delta_4 y$		$\Delta_4 c$	$\Delta_4 h$	$\Delta_4 y$
Mean (%)		2.077	2.522	2.143		1.456	-0.858	1.755
	$\Delta_4 c$	1.237			$\Delta_4 c$	1.563		
St. dev (%)/Corr.	$\Delta_4 h$	0.396	5.417		$\Delta_4 h$	0.582	2.686	
	$\Delta_4 y$	0.728	0.254	1.253	$\Delta_4 y$	0.618	0.434	1.634
		Italy				**Spain**		
		$\Delta_4 c$	$\Delta_4 h$	$\Delta_4 y$		$\Delta_4 c$	$\Delta_4 h$	$\Delta_4 y$
Mean (%)		2.129	1.860	1.919		2.755	4.790	2.969
	$\Delta_4 c$	2.016			$\Delta_4 c$	2.048		
St. dev (%)/Corr.	$\Delta_4 h$	0.106	8.889		$\Delta_4 h$	0.719	8.928	
	$\Delta_4 y$	0.776	-0.016	1.622	$\Delta_4 y$	0.839	0.642	1.694
		UK				**US**		
		$\Delta_4 c$	$\Delta_4 h$	$\Delta_4 y$		$\Delta_4 c$	$\Delta_4 h$	$\Delta_4 y$
Mean (%)		2.894	4.782	2.368		3.168	1.335	2.883
	$\Delta_4 c$	2.168			$\Delta_4 c$	1.489		
St. dev (%)/Corr.	$\Delta_4 h$	0.712	8.059		$\Delta_4 h$	0.470	3.274	
	$\Delta_4 y$	0.816	0.567	1.859	$\Delta_4 y$	0.798	0.318	1.882

For each country, the first row of the table shows the mean (in percentage points) of the yearly rates of growth of private final consumption expenditure, real house prices and GDP. The Standard deviation/Correlation matrices report standard deviations of the growth rates (in percentage points) on the diagonal and contemporaneous correlation coefficients below the diagonal. Sample period: 1979(1)-2007(4), except for Spain (1980(1)-2007(4)) and Italy (1979(1)-2007(3)).

Table 2
Cointegration analysis

	France		**Germany**		**Italy**	
Eigenvalues						
	0.208		0.096		0.184	
	0.037		0.014		0.042	
	0.026		0.000		0.022	
H_0: rank $\leq$ (p-value)						
0	0.015		0.865		0.075	
1	0.548		0.998		0.515	
2	0.085		0.959		0.110	
	Unrestricted		Unrestricted		Unrestricted	
	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$
c	1	0.004	1	-0.025	1	-0.134
		(0.058)		(0.056)		(0.039)
h	-0.129	0.282	-0.312	0.025	0.012	0.347
	(0.016)	(0.080)	(0.078)	(0.031)	(0.023)	(0.154)
y	-0.837	0.098	-1.054	0.112	-1.005	-0.001
	(0.015)	(0.038)	(0.040)	(0.047)	(0.024)	(0.037)
BIC	-21.242		-20.421		-19.005	
AIC	-22.459		-21.201		-19.847	
	Restricted		Restricted		Restricted	
	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$
c	1	0	1	0	1	-0.129
						(0.037)
h	-0.129	0.283	-0.256	0	0	0.375
	(0.016)	(0.080)	(0.067)			(0.154)
y	-0.837	0.097	-1	0.122	-1	0
	(0.015)	(0.032)		(0.040)		
LRT (p-value)	0.946		0.468		0.970	
BIC	-21.452		-20.629		-19.233	
AIC	-22.527		-21.267		-19.882	

(continued)

(Table 2 continued)

	Spain		**UK**		**US**	
Eigenvalues						
	0.173		0.102		0.189	
	0.044		0.056		0.042	
	0.002		0.023		0.022	
H_0: rank $\leq$ (p-value)						
0	0.128		0.306		0.010	
1	0.793		0.334		0.211	
2	0.661		0.102		0.105	
	Unrestricted		Unrestricted		Unrestricted	
	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$
c	1	-0.150	1	-0.056	1	0.077
		(0.034)		(0.040)		(0.047)
h	-0.001	-0.227	-0.001	0.012	-0.161	0.264
	(0.020)	(0.134)	(0.049)	(0.069)	(0.029)	(0.064)
y	-0.981	-0.063	-1.127	0.043	-1.008	0.121
	(0.036)	(0.047)	(0.089)	(0.031)	(0.014)	(0.051)
BIC	-19.883		-18.538		-21.219	
AIC	-20.472		-19.535		-22.217	
	Restricted		Restricted		Restricted	
	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$	$\boldsymbol{\beta}$	$\boldsymbol{\alpha}$
c	1	-0.110	1	-0.087	1	0
		(0.028)		(0.027)		
h	0	0	0	0	-0.158	0.235
					(0.023)	(0.060)
y	-1	0	-1.120	0	-1	0.068
			(0.035)			(0.042)
LRT (p-value)	0.180		0.570		0.249	
BIC	-20.037		-18.709		-21.369	
AIC	-20.479		-19.564		-22.224	

Table 3
Long-run effects of permanent shocks

	France		**Germany**		**Italy**	
Variable	$\boldsymbol{\psi}_1$	$\boldsymbol{\psi}_2$	$\boldsymbol{\psi}_1$	$\boldsymbol{\psi}_2$	$\boldsymbol{\psi}_1$	$\boldsymbol{\psi}_2$
c	0.673	0.538	0.714	0.424	0.711	0
	(0.318)	(0.187)	(0.229)	(0.135)	(0.152)	
h	0.976	4.173	0.709	1.658	2.190	4.138
	(1.876)	(1.452)	(0.655)	(0.528)	(1.450)	(0.859)
y	0.653	0	0.532	0	0.711	0
	(0.169)		(0.092)		(0.152)	
	Spain		**UK**		**US**	
Variable	$\boldsymbol{\psi}_1$	$\boldsymbol{\psi}_2$	$\boldsymbol{\psi}_1$	$\boldsymbol{\psi}_2$	$\boldsymbol{\psi}_1$	$\boldsymbol{\psi}_2$
c	1.126	0	1.326	0	0.853	0.367
	(0.284)		(0.368)		(0.263)	(0.113)
h	4.425	3.741	4.567	2.456	0.327	2.323
	(1.702)	(0.834)	(1.560)	(0.646)	(0.924)	(0.713)
y	1.126	0	1.184	0	0.802	0
	(0.284)		(0.328)		(0.198)	

This table reports, for each country, the estimated elements of matrx $\mathbf{A}$ in the common trends structural representation. Asymptotic standard errors are in parentheses.

Table 4
Transitory component: descriptive statistics

	France				**Germany**			
		c^{TR}	h^{TR}	y^{TR}		c^{TR}	h^{TR}	y^{TR}
	c^{TR}	1.330			c^{TR}	0.897		
St. dev (%)/Corr.	h^{TR}	0.967	11.26		h^{TR}	0.866	2.216	
	y^{TR}	0.825	0.723	1.440	y^{TR}	-0.299	-0.155	1.423
	Italy				**Spain**			
		c^{TR}	h^{TR}	y^{TR}		c^{TR}	h^{TR}	y^{TR}
	c^{TR}	1.291			c^{TR}	4.362		
St. dev (%)/Corr.	h^{TR}	0.324	6.786		h^{TR}	0.982	15.167	
	y^{TR}	0.246	0.036	0.364	y^{TR}	0.992	0.993	2.882
	UK				**US**			
		c^{TR}	h^{TR}	y^{TR}		c^{TR}	h^{TR}	y^{TR}
	c^{TR}	1.977			c^{TR}	1.368		
St. dev (%)/Corr.	h^{TR}	0.790	8.144		h^{TR}	0.886	5.637	
	y^{TR}	0.254	0.627	1.173	y^{TR}	0.985	0.899	1.830

For each country, the Standard deviation/Correlation matrices report standard deviations (in percentage points) on the diagonal and contemporaneous correlation coefficients below the diagonal.

Table 5
Forecast error variance decomposition

Variable	Horizon (quarters)	France ψ^1	France ψ^2	France ν	Germany ψ^1	Germany ψ^2	Germany ν	Italy ψ^1	Italy ψ^2	Italy ν
c	1	0.658	0.337	0.006	0.976	0.001	0.024	0.302	0.141	0.557
	4	0.752	0.185	0.063	0.944	0.043	0.013	0.648	0.090	0.263
	12	0.652	0.280	0.068	0.847	0.148	0.005	0.845	0.044	0.111
	20	0.577	0.387	0.035	0.808	0.189	0.003	0.901	0.028	0.072
	∞	0.610 (0.264)	0.390 (0.264)	0	0.739 (0.163)	0.261 (0.163)	0	1	0	0
h	1	0.240	0.401	0.359	0.224	0.774	0.002	0.040	0.925	0.034
	4	0.150	0.502	0.348	0.161	0.838	0.001	0.019	0.974	0.007
	12	0.038	0.765	0.197	0.157	0.842	0.001	0.138	0.852	0.010
	20	0.023	0.893	0.084	0.157	0.843	0.001	0.179	0.815	0.007
	∞	0.052 (0.187)	0.948 (0.187)	0	0.155 (0.241)	0.845 (0.241)	0	0.219 (0.218)	0.781 (0.218)	0
y	1	0.590	0.051	0.359	0.470	0.012	0.518	0.900	0.027	0.073
	4	0.593	0.123	0.284	0.520	0.010	0.471	0.973	0.016	0.011
	12	0.732	0.065	0.203	0.703	0.064	0.233	0.992	0.005	0.003
	20	0.837	0.064	0.099	0.786	0.056	0.159	0.995	0.003	0.002
	∞	1	0	0	1	0	0	1	0	0

Variable	Horizon (quarters)	Spain ψ^1	Spain ψ^2	Spain ν	UK ψ^1	UK ψ^2	UK ν	US ψ^1	US ψ^2	US ν
c	1	0.000	0.058	0.942	0.652	0.004	0.344	0.898	0.022	0.080
	4	0.124	0.013	0.863	0.761	0.082	0.156	0.826	0.015	0.160
	12	0.516	0.023	0.461	0.904	0.040	0.056	0.830	0.111	0.059
	20	0.733	0.016	0.251	0.943	0.023	0.034	0.821	0.149	0.030
	∞	1	0	0	1	0	0	0.844 (0.113)	0.156 (0.113)	0
h	1	0.051	0.857	0.092	0.147	0.841	0.012	0.000	0.237	0.763
	4	0.117	0.753	0.130	0.250	0.739	0.011	0.018	0.459	0.523
	12	0.251	0.649	0.099	0.577	0.410	0.013	0.005	0.797	0.198
	20	0.351	0.588	0.062	0.667	0.326	0.008	0.010	0.910	0.081
	∞	0.583 (0.214)	0.417 (0.214)	0	0.776 (0.151)	0.224 (0.151)	0	0.019 (0.108)	0.981 (0.108)	0
y	1	0.794	0.168	0.038	0.937	0.027	0.037	0.124	0.708	0.168
	4	0.744	0.057	0.198	0.979	0.006	0.015	0.368	0.402	0.230
	12	0.836	0.017	0.148	0.994	0.002	0.004	0.711	0.163	0.127
	20	0.906	0.010	0.084	0.995	0.001	0.004	0.836	0.093	0.071
	∞	1	0	0	1	0	0	1	0	0

Table 6
Consumption elasticities to real house price changes due to the permanent shock ψ^2

Horizon (quarters)	4	8	12	16	20	∞
France	0.13 *	0.14 *	0.14 *	0.13 *	0.13 *	0.13 *
Germany	0.18 *	0.23 *	0.24 *	0.25 *	0.25 *	0.26 *
US	0.14 *	0.19 *	0.19 *	0.18 *	0.17 *	0.16 *
Italy	0.06 *	0.03 *	0.01 *	0.01	0.00	0
Spain	0.01 *	0.04 *	0.03 *	0.03 *	0.02 *	0
UK	0.12 *	0.07 *	0.04	0.02	0.02	0

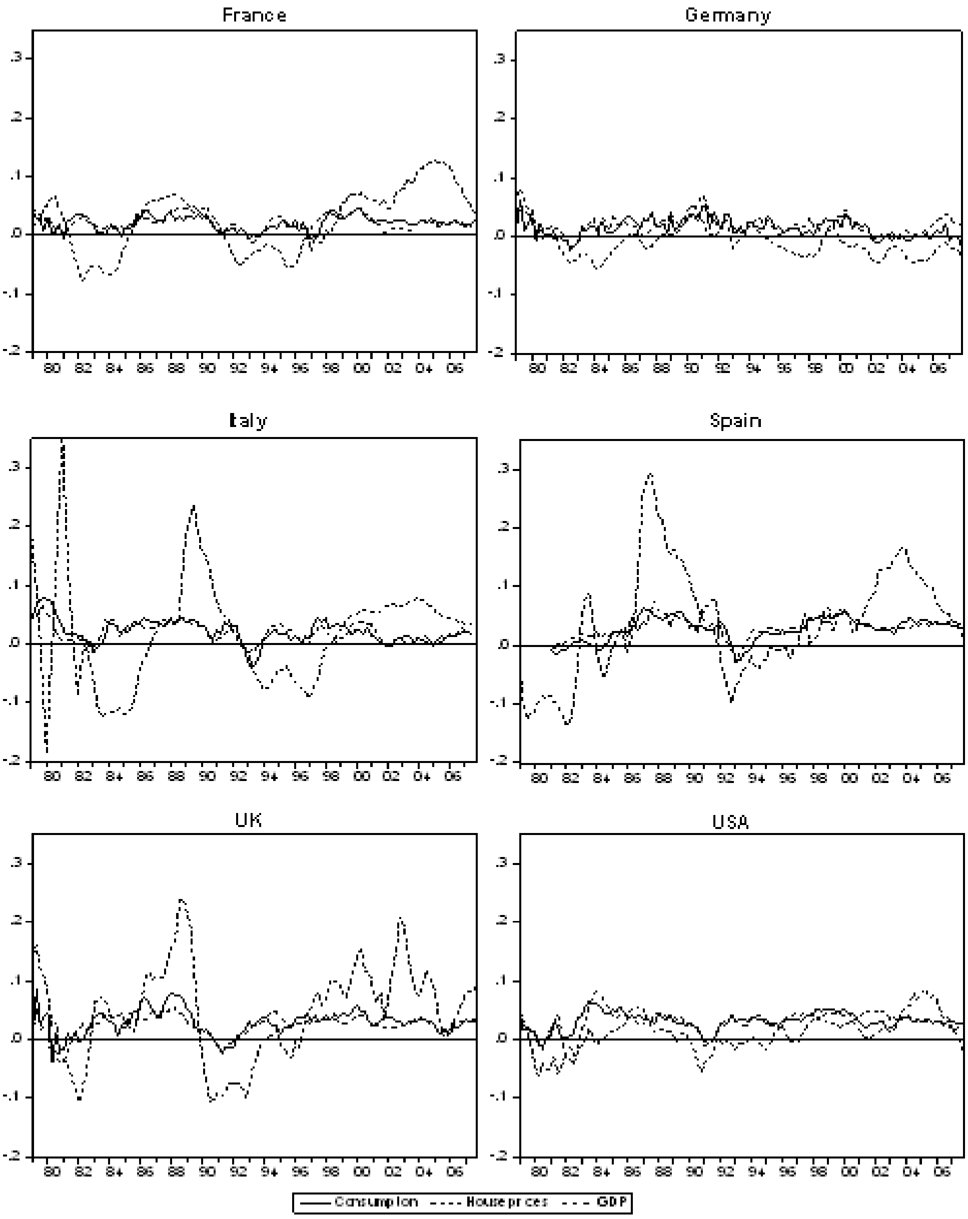
The table reports the percentage response of consumption expenditure to a 1% change in real house prices due to the output-neutral permanent shock ψ^2. * denotes statistical significance at the 5% level.

Table 7
Principal components analysis of the permanent and transitory components of real house prices

	PC_1	PC_2	PC_3	PC_4	PC_5	PC_6
$\hat{h}^P$	0.815	0.107	0.047	0.018	0.009	0.004
$\hat{h}^P_{FR}$	0.896	0.023	0.019	0.051	0.009	0.002
$\hat{h}^P_{GE}$	0.787	0.120	0.043	0.048	0.001	0.001
$\hat{h}^P_{IT}$	0.485	0.472	0.041	0.002	0.000	0.000
$\hat{h}^P_{SP}$	0.812	0.015	0.166	0.003	0.001	0.005
$\hat{h}^P_{UK}$	0.964	0.002	0.014	0.000	0.002	0.017
$\hat{h}^P_{US}$	0.946	0.009	0.001	0.003	0.041	0.000
$\hat{h}^{\tau_1}$	0.993	0.004	0.001	0.001	0.000	0.000
$\hat{h}^{\tau_1}_{FR}$	0.998	0.000	0.000	0.001	0.000	0.001
$\hat{h}^{\tau_1}_{GE}$	0.992	0.005	0.002	0.000	0.001	0.000
$\hat{h}^{\tau_1}_{IT}$	0.987	0.010	0.003	0.000	0.000	0.000
$\hat{h}^{\tau_1}_{SP}$	0.993	0.004	0.001	0.001	0.000	0.001
$\hat{h}^{\tau_1}_{UK}$	0.991	0.006	0.000	0.001	0.000	0.000
$\hat{h}^{\tau_1}_{US}$	0.996	0.000	0.002	0.001	0.001	0.000
$\hat{h}^{\tau_2}$	0.697	0.264	0.022	0.011	0.004	0.002
$\hat{h}^{\tau_2}_{FR}$	0.345	0.592	0.056	0.007	0.000	0.000
$\hat{h}^{\tau_2}_{GE}$	0.919	0.066	0.000	0.002	0.008	0.006
$\hat{h}^{\tau_2}_{IT}$	0.916	0.013	0.046	0.024	0.002	0.000
$\hat{h}^{\tau_2}_{SP}$	0.947	0.020	0.002	0.020	0.012	0.000
$\hat{h}^{\tau_2}_{UK}$	0.978	0.007	0.000	0.003	0.004	0.008
$\hat{h}^{\tau_2}_{US}$	0.076	0.883	0.028	0.013	0.000	0.000
$\hat{h}^{TR}$	0.389	0.219	0.155	0.124	0.063	0.050
$\hat{h}^{TR}_{FR}$	0.714	0.010	0.013	0.029	0.214	0.020
$\hat{h}^{TR}_{GE}$	0.191	0.086	0.453	0.268	0.002	0.000
$\hat{h}^{TR}_{IT}$	0.037	0.815	0.000	0.056	0.028	0.064
$\hat{h}^{TR}_{SP}$	0.572	0.018	0.183	0.104	0.007	0.115
$\hat{h}^{TR}_{UK}$	0.379	0.058	0.220	0.283	0.035	0.025
$\hat{h}^{TR}_{US}$	0.440	0.329	0.061	0.005	0.092	0.074

The table reports the results of the principal components (PC) analysis conducted on 6 sub-sets of (standardized) series, each comprising the same variable for the 6 countries investigated. For each set the first row shows the fraction of the total variance explained by each PC_i $(i = 1, ...6)$; the subsequent six rows display the fraction of the variance of the individual series attributable to each PC_i.

Figure 1
Consumption, house prices and GDP



The figure displays yearly rates of change of private final consumption expenditure, real house prices and GDP for France, Germany, Italy, Spain, the UK,and the US over the period 1979-2007.

Figure 2
Cointegrating vectors

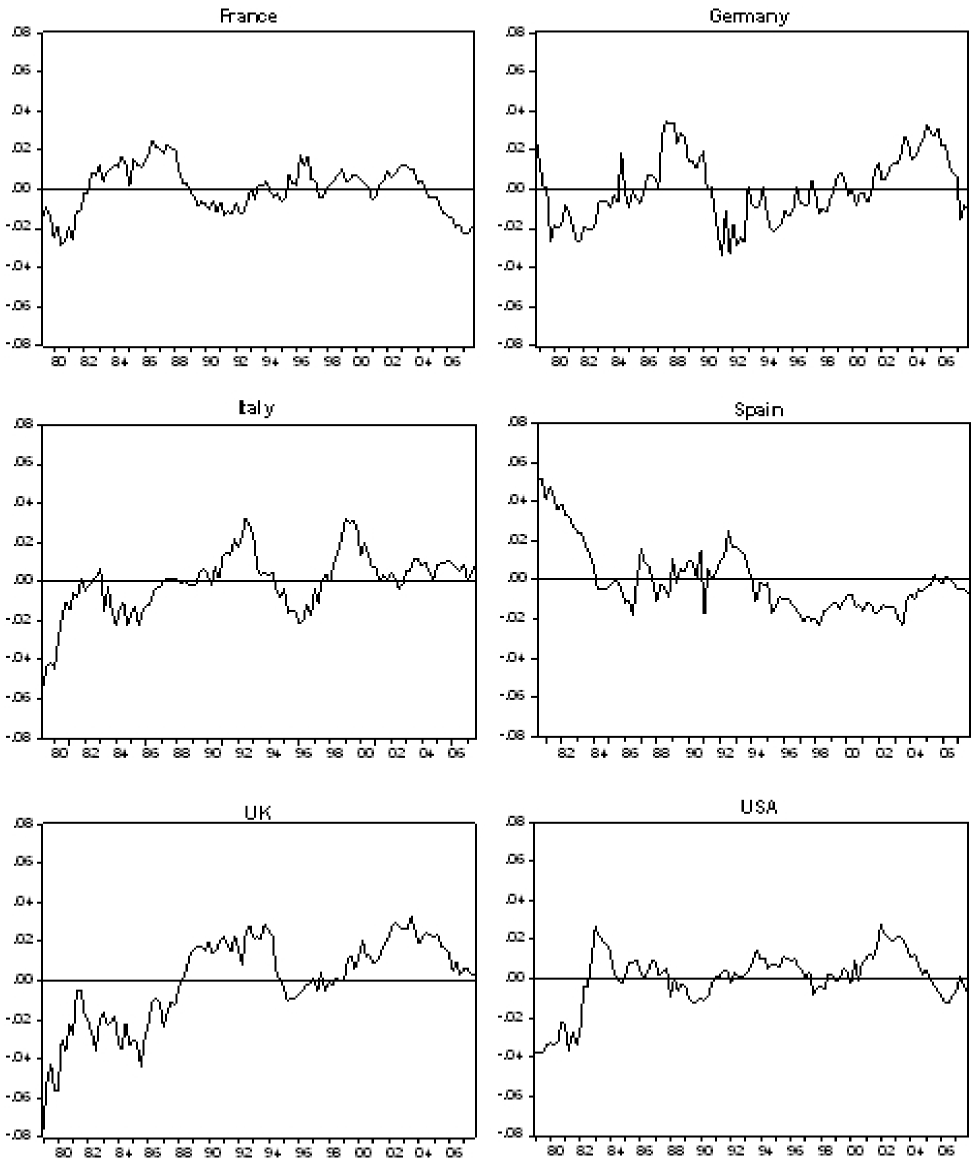


Figure 3
Transitory components of consumption, house prices and GDP

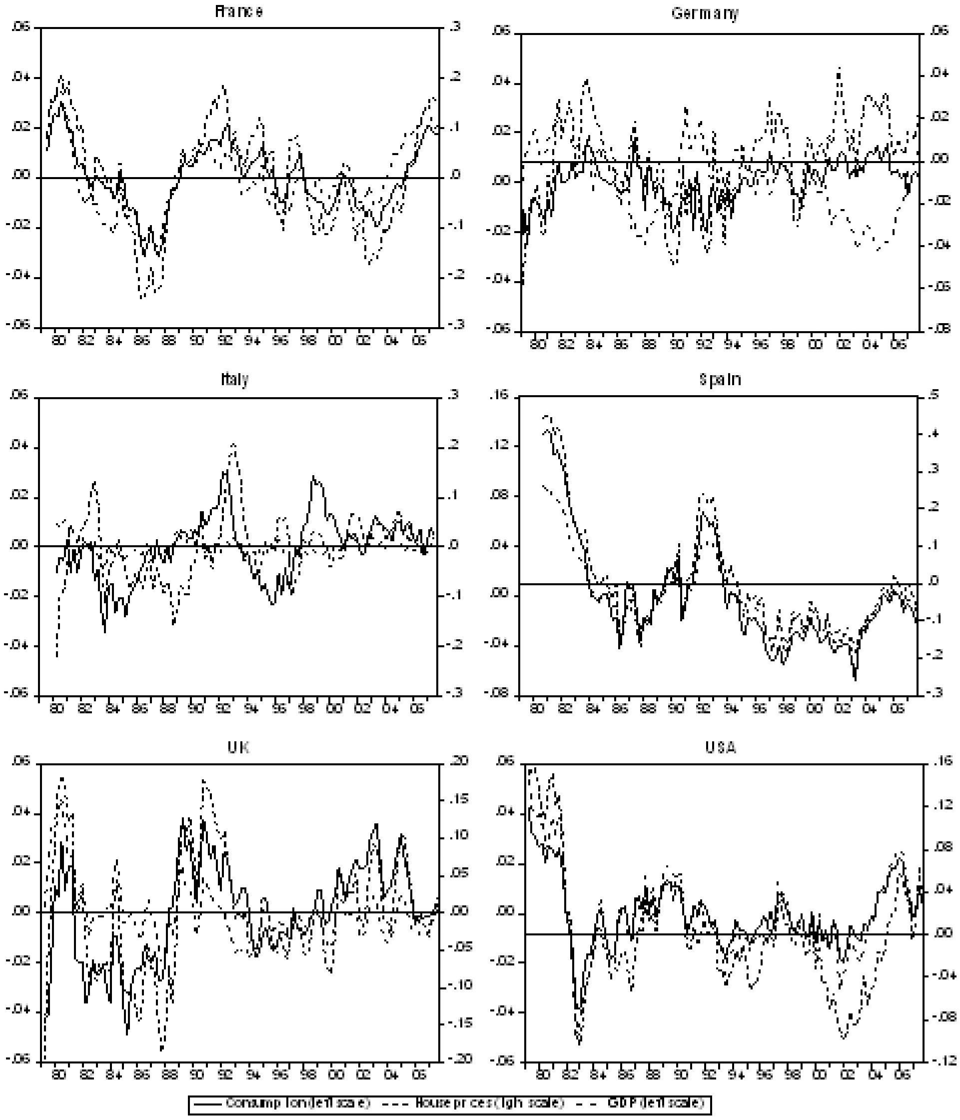
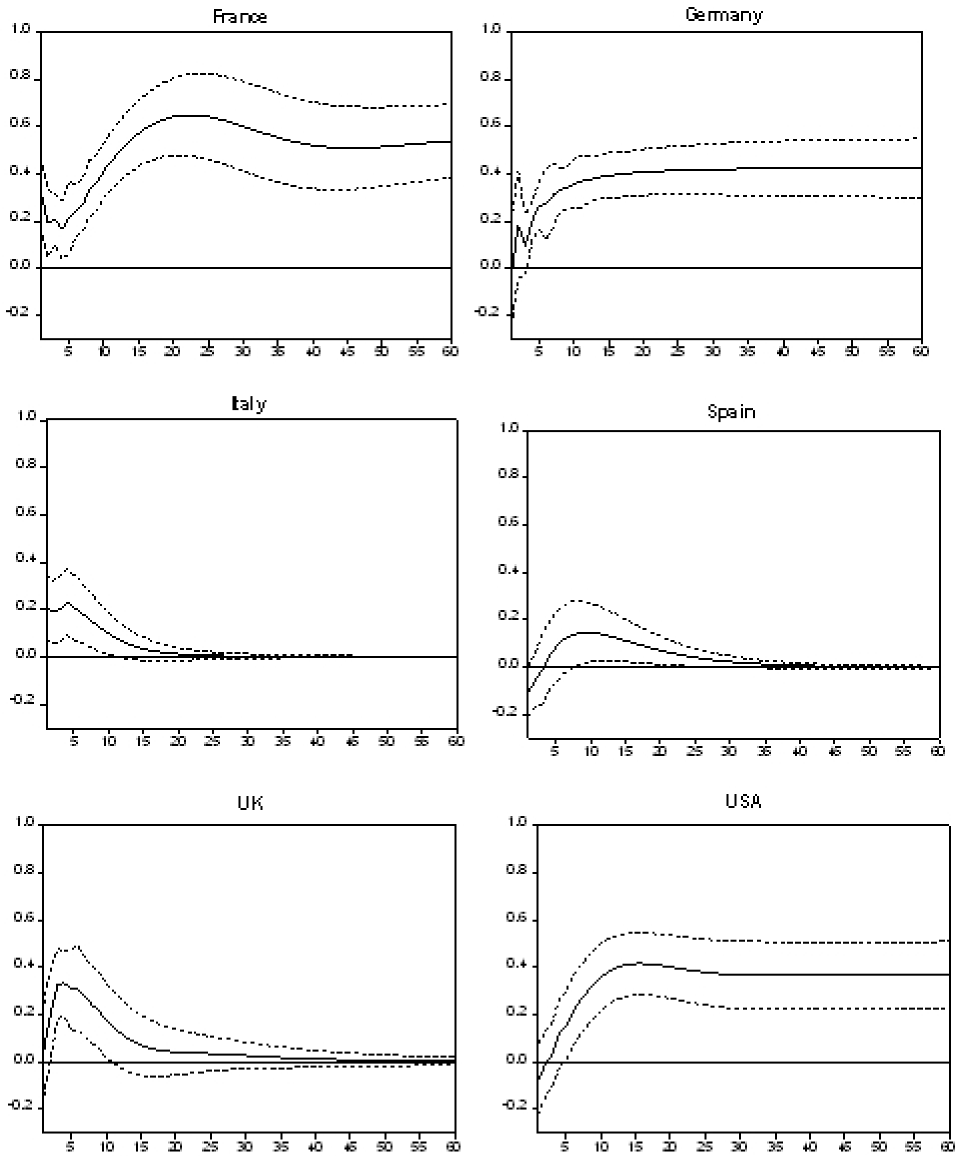
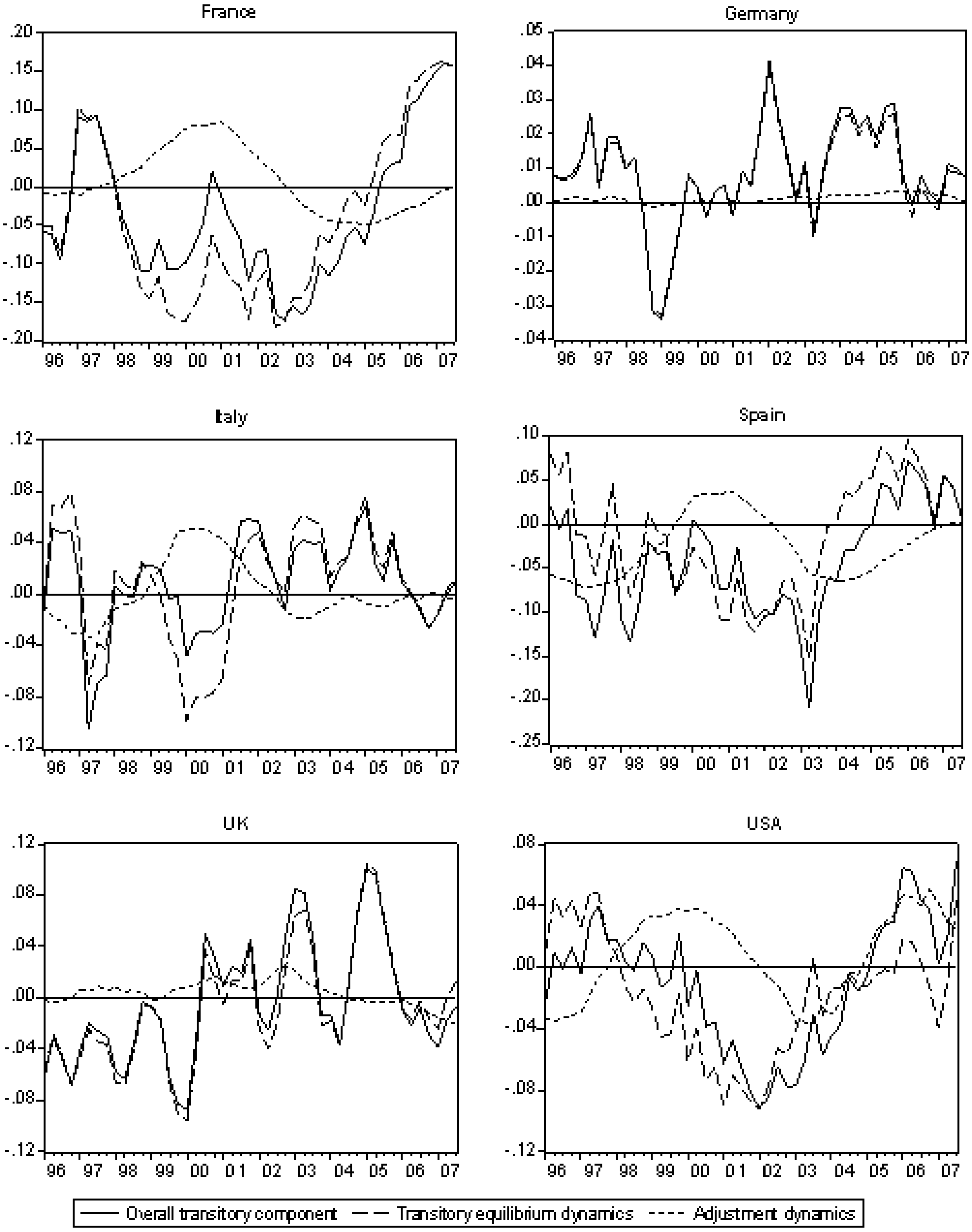


Figure 4
Consumption response to a permanent shock ψ^2



The figure displays, for each country, the response over 60 quarters of private final consumption to a unitary permanent shock ψ^2 with one standard deviation bands.

Figure 5
Analysis of transitory house price component: 1996-2007



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