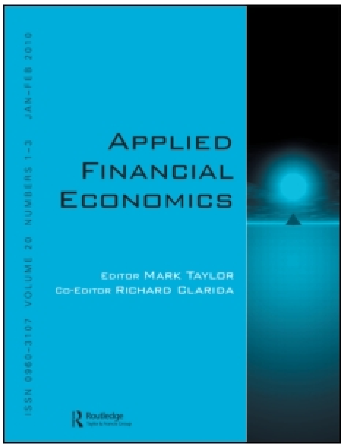

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Permanent and transitory dynamics in house prices and consumption: some implications for the real effects of the financial crisis

Fabio C. Bagliano [ab]; Claudio Morana [bcd]
[a] Dipartimento di Scienze Economiche e Finanziarie, Università di Torino, Torino, Italy [b] CeRP, Moncalieri, Italy [c] Dipartimento di Scienze Economiche e Metodi Quantitativi, Università del Piemonte Orientale, Novara, Italy [d] International Centre for Economic Research (ICER), Torino, Italy

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Permanent and transitory dynamics in house prices and consumption: some implications for the real effects of the financial crisis

Fabio C. Bagliano[a,b,]* and Claudio Morana[b,c,d]

[a]*Dipartimento di Scienze Economiche e Finanziarie, Università di Torino, Torino, Italy*
[b]*CeRP, Collegio Carlo Alberto, Moncalieri, Italy*
[c]*Dipartimento di Scienze Economiche e Metodi Quantitativi, Università del Piemonte Orientale, Novara, Italy*
[d]*International Centre for Economic Research (ICER), Torino, Italy*

In this article, 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, interactions between house prices and consumption are detected in all countries at shorter horizons, with important implications of the current unwinding of the sub-prime crisis for real activity. Evidence for international comovements in the common trend component of house price dynamics is also found.

I. Introduction

The aim of this article 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 *et al*. (1991) to investigate the impact of permanent and temporary real house price dynamics on consumption using quarterly data over the period 1979 to 2007 for the major European economies (France, Germany, Italy, Spain and the

*Corresponding author. E-mail: fabio.bagliano@unito.it

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UK) and the US.[1] 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 detected (Italy, Spain and the UK), house prices and consumption appear to be related over the short- to medium-term horizon. On the basis of the estimated elasticities, a sizeable medium-term impact on the consumption of a decline in real house prices of the magnitude experienced by the US in 2008 can be predicted for various countries (reaching −4.6% and −6% at the peak for the US and Germany, respectively, and ranging between −1% and −3.4% for the other countries). Though subject to several caveats (discussed below), those results may provide an useful benchmark measure of one important dimension of the real impact of the ongoing financial crisis.

Finally, a principal component analysis carried out on the permanent component of real house price dynamics provides a strong evidence for global comovements across countries.

This article is organized as follows. In the next section, the econometric methodology is outlined and the data set is described. Section III presents the empirical results, concerning 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 IV explores some implications for the ongoing crisis in housing markets. A final section summarizes the main conclusions.

II. 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 *et al.* (1991) and Mellander *et al.* (1992), exploiting 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 analysed 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.

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 Gross Domestic Product (GDP)). If there exist $0<r<n$ cointegrating relations among the variables, the following cointegrated Vector Autoregression (VAR) representation for $\mathbf{x}_t$ holds (deterministic terms are omitted throughout for ease of exposition)

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

where $\mathbf{\Pi}(L) = \mathbf{\Pi}_1 + \mathbf{\Pi}_2 L + \cdots + \mathbf{\Pi}_p L^{p-1}$ 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 *et al.* (1992), the cointegrated VAR in Equation 1 can be

[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.

inverted to yield the following stationary Wold representation for $\Delta \mathbf{x}_t$:

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

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + \cdots$ with $\sum_{j=0}^{\infty} j\,|\mathbf{C}_j| < \infty$. Starting from Equation 2 it is possible to derive the *stochastic trends* representation of $\mathbf{x}_t$, decomposing the series into a permanent (nonstationary) 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 ε 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 nonstationary component of $\mathbf{x}$ can be expressed in terms of a reduced number $k = n - r$ of *common stochastic trends* as follows:

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

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 Equation 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 nonstationary endogenous variables (the transitory component $\mathbf{C}^*(L)\boldsymbol{\varepsilon}_t$ being stationary). Hence,

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

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 Equations 2 and 3, the vector of reduced form disturbances ε must be transformed into a vector of underlying structural shocks, some with *permanent* effects on the level of $\mathbf{x}$ and some with only *transitory* effects. Let us denote this vector of independent and identically distributed (i.i.d.) structural disturbances as $\boldsymbol{\varphi}_t \equiv \binom{\boldsymbol{\psi}_t}{\boldsymbol{\nu}_t}$, 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 $\boldsymbol{\Gamma}(L) = \boldsymbol{\Gamma}_0 + \boldsymbol{\Gamma}_1 L + \cdots$. Since the first element of $\mathbf{C}(L)$ in Equation 2 is $\mathbf{I}$, equating the first term of the right-hand sides of Equations 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 $\boldsymbol{\Gamma}_0$ is an invertible matrix. Hence, the comparison of Equations 6 and 2 shows that

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

implying that $\mathbf{C}(1)\boldsymbol{\Gamma}_0 = \boldsymbol{\Gamma}(1)$. 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 $\boldsymbol{\Gamma}(1)$ must be imposed:

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

with $\mathbf{A}$ being the $n \times k$ matrix in Equation 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 $\boldsymbol{v}_t$ are restricted to have only transitory effects.

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

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

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

As shown in detail by Stock and Watson (1988), King *et al.* (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 Equation 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 Equations 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 $E(\boldsymbol{\psi}_t\boldsymbol{\psi}_t') = \mathbf{I}$)

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

where $\boldsymbol{\Sigma}$ is the variance/covariance matrix of the VAR innovations $\boldsymbol{\varepsilon}$. The remaining $k(k-1)/2$ restrictions needed for (exact) the identification of $\mathbf{A}$ then have 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 the identification of $\mathbf{A}$ is achieved, estimates of the structural permanent disturbances may be obtained from Equation 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 Equation 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 permanent ($\boldsymbol{\psi}_t$) and transitory shocks ($\boldsymbol{v}_t$). From Equation 9

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

where the first component $\boldsymbol{\Gamma}_{\psi}^*(L)\boldsymbol{\psi}_t$ gives the contribution of permanent innovations to the overall transitory fluctuations (dynamics 'along the attractor'), while the vector $\boldsymbol{\Gamma}_2^*(L)\boldsymbol{v}_t$ 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.

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 Organization for Economic Co-operation and Development (OECD); in particular, house price data are extensively described and analysed in Girouard *et al.* (2006). The sample period ranges from 1978(1)

[2] In addition, nr restrictions are needed to identify r separate transitory shocks in $\boldsymbol{v}_t$. To this aim, a set of $kr + r(r+1)/2$ restrictions are provided by the orthogonality conditions $E[\boldsymbol{\psi}_t, \boldsymbol{v}_t'] = \mathbf{0}$ and $E[\boldsymbol{v}_t, \boldsymbol{v}_t'] = \mathbf{I}_r$. 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.

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 International Monetary Fund (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 SDs 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, 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.

III. 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.

Cointegration analysis

To test for the existence of long-term relationships among the investigated variables, different criteria have been jointly employed. The standard 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 behaviour 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 Akaike Information Criterion (AIC) and Bayesian Information Criterion (BIC) 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 the 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 Fig. 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

[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 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.

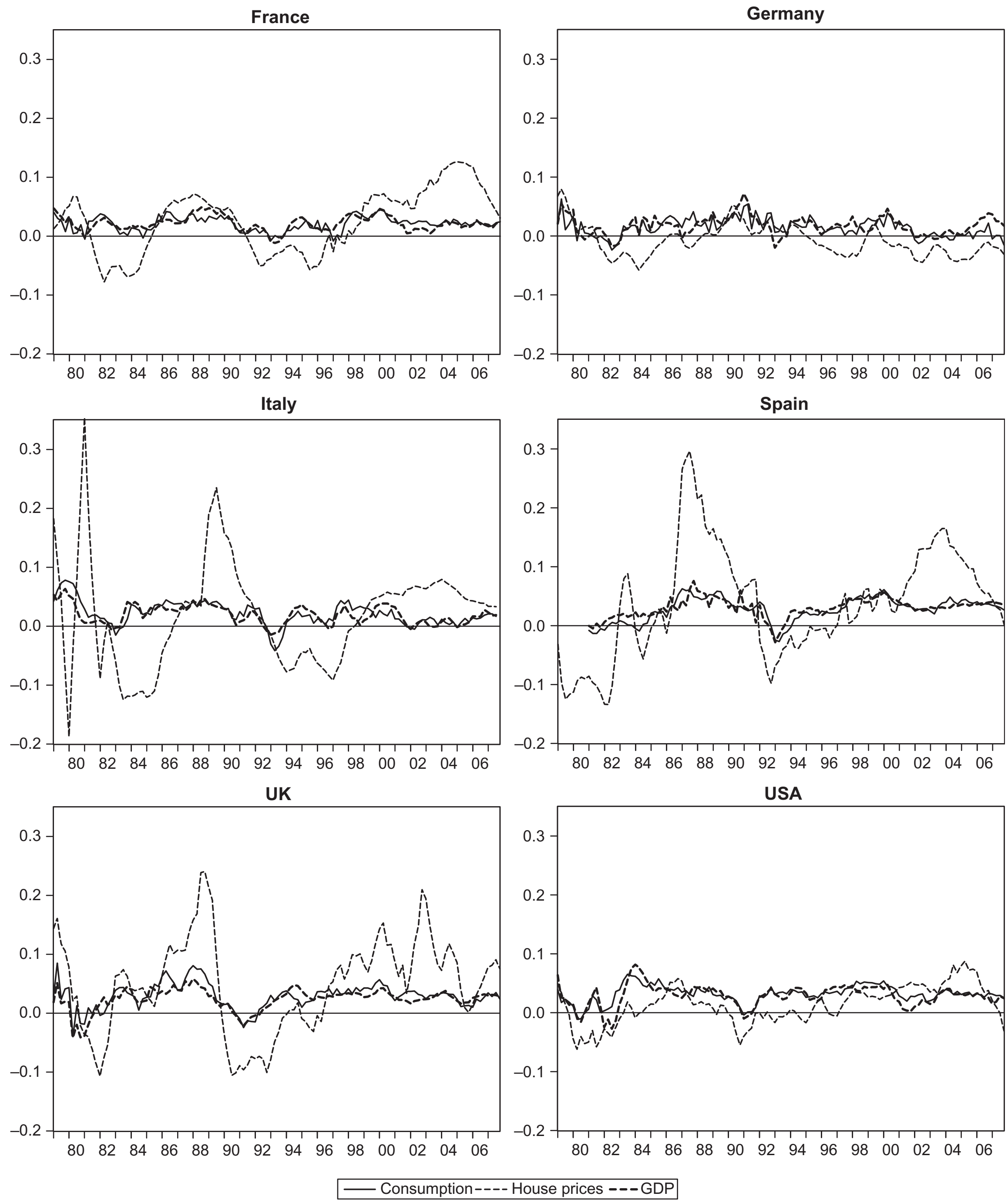

Fig. 1. Consumption, house prices and GDP
Note: The figure displays yearly rates of change of private final consumption expenditure, real house prices and GDP for France, Germany, Italy, Spain, the UK and the US over the period 1979–2007.

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

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

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
SD (%)/corr.	$\Delta_4 c$	1.237			$\Delta_4 c$	1.563		
	$\Delta_4 h$	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
SD (%)/corr.	$\Delta_4 c$	2.016			$\Delta_4 c$	2.048		
	$\Delta_4 h$	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
SD (%)/corr.	$\Delta_4 c$	2.168			$\Delta_4 c$	1.489		
	$\Delta_4 h$	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

Notes: In the following three rows, data are organized in 3 by 3 matrices, with SDs of the growth rates (SD, in percentage points) on the diagonal and contemporaneous correlation coefficients (corr.) below the diagonal. Sample period: January 1979– April 2007, except for Spain (January 1980–April 2007) and Italy (January 1979–March 2007).

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	
	β	α	β	α	β	α
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	
	β	α	β	α	β	α
c	1	0	1	0	1	−0.129
						(0.037)

(*continued*)

Table 2. Continued

	France		Germany		Italy	
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	
	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 ≤ (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	
	β	α	β	α	β	α
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	
	β	α	β	α		
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

Note: LRT = Likelihood Ratio Test.

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 collateralized 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 *et al.* (2006) points to a gradually weakening

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

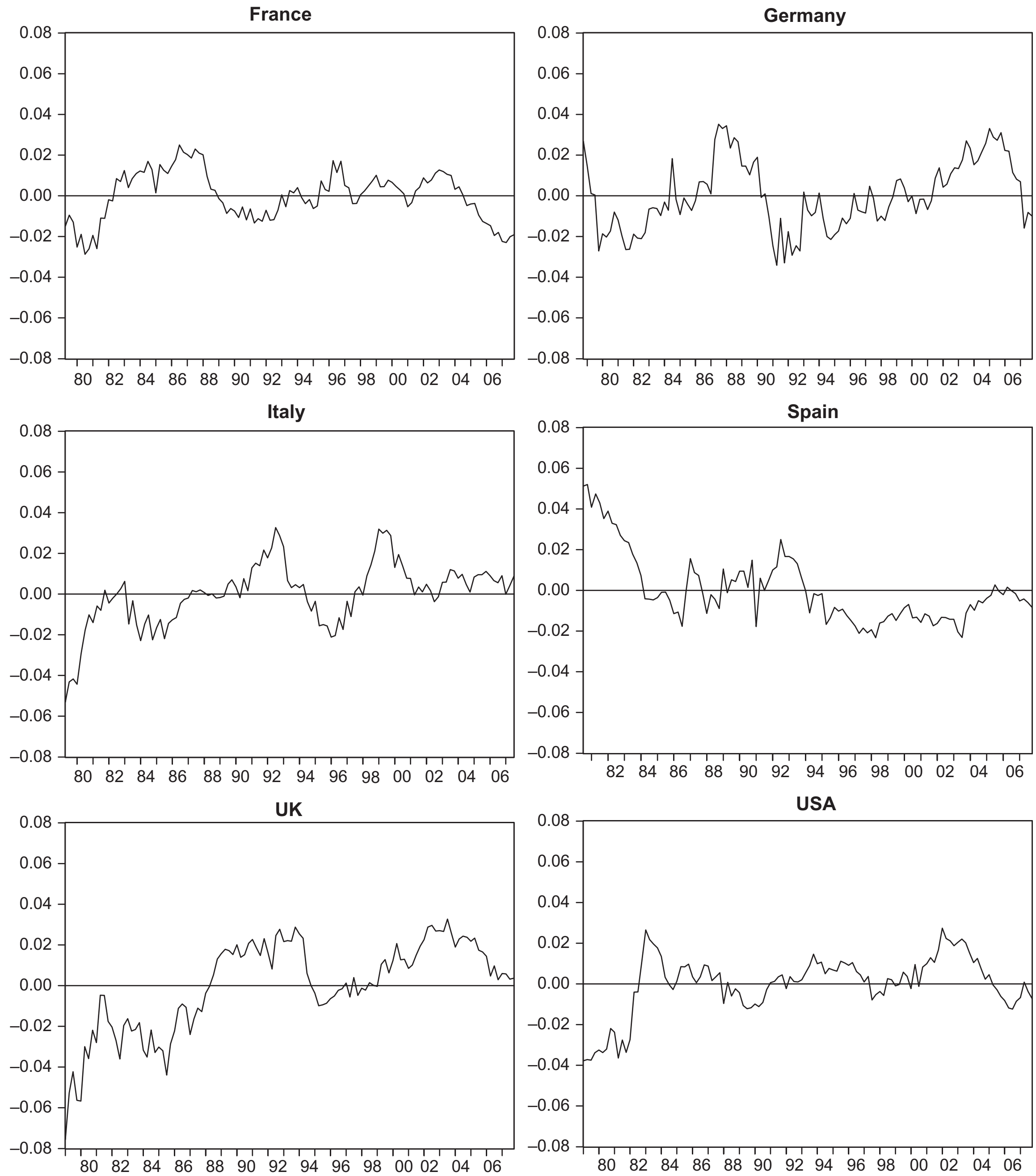

Fig. 2. Cointegrating vectors

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 coefficients in α 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).

Permanent and transitory components

The existence of one cointegrating relationship among three $I(1)$ nonstationary variables implies the presence of two distinct sources of shocks having permanent effects on at least some of the variables, and one transitory shock.[7] In terms of the common stochastic trend representation in Equation 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:

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

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 Equations 10 and 12, derived from the (restricted 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. 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 therefore 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, yielding 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 Fig. 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

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

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

Table 3. Long-run effects of permanent shocks

Variable	ψ_1	ψ_2	ψ_1	ψ_2	ψ_1	ψ_2
	France		Germany		Italy	
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	
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)	

Notes: This table reports, for each country, the estimated elements of matrix **A** in the common trends structural representation. Asymptotic SDs are in parentheses.

Table 4. Transitory component: descriptive statistics

SD (%)/corr.		c^{TR}	h^{TR}	y^{TR}		c^{TR}	h^{TR}	y^{TR}
	France				Germany			
	c^{TR}	1.330			c^{TR}	0.897		
	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}	1.291			c^{TR}	4.362		
	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}	1.977			c^{TR}	1.368		
	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

Note: For each country, data are organized in 3 by 3 matrices, with SDs (SD, in percentage points) on the diagonal and contemporaneous correlation coefficients (corr.) below the diagonal.

only with respect to consumption, whereas in the case of Italy both correlations are only around 0.3.

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 decomposition exercise has been carried out at different horizons, including the business cycle range (1–5 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 1-year horizon. The output-neutral shock ψ^2 accounts for the bulk of house price

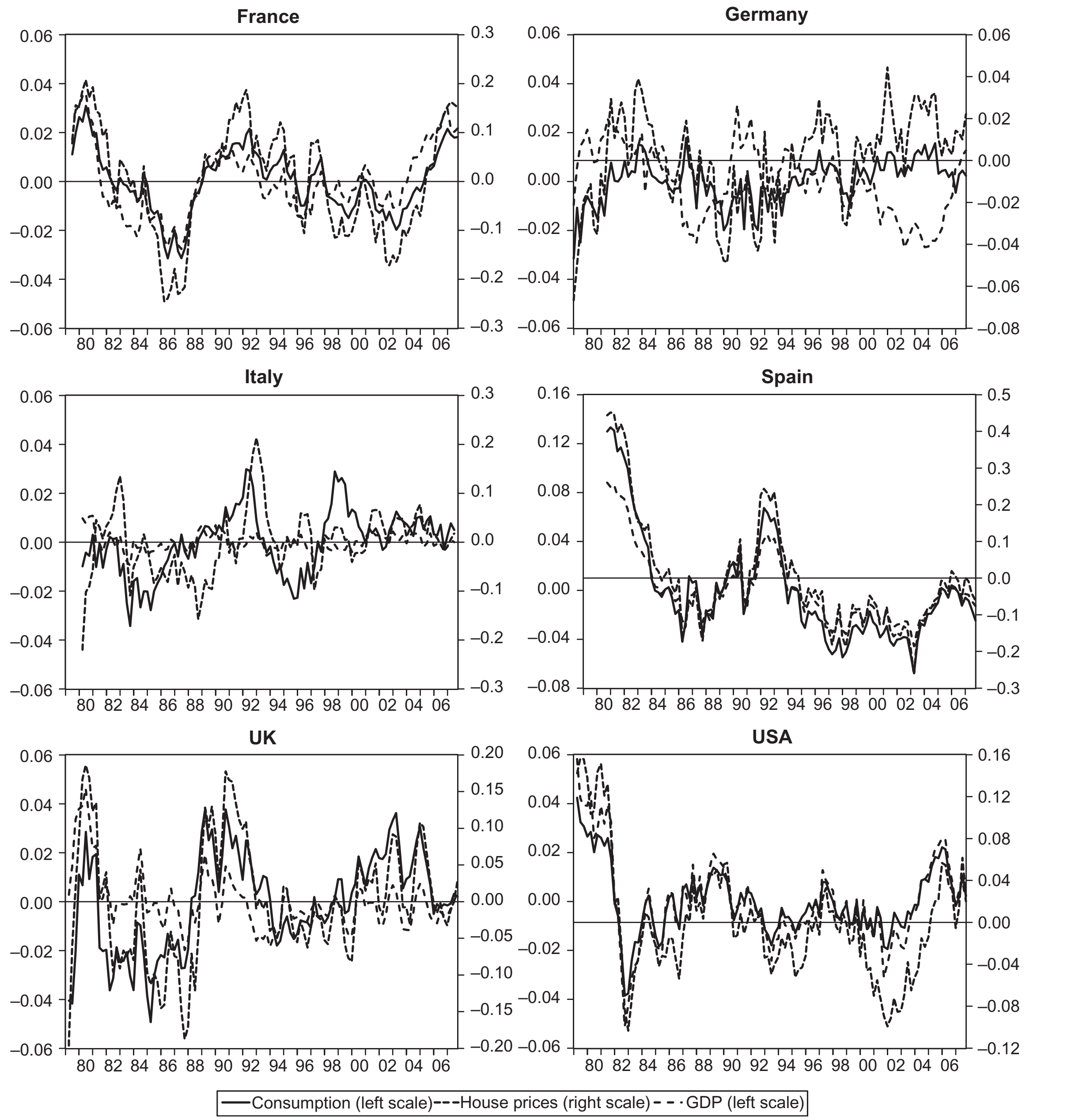

Fig. 3. Transitory components of consumption, house prices and GDP

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 nonnegligible 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 SD 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

Table 5. Forecast error variance decomposition

Variable	Horizon (quarters)	ψ^1	ψ^2	ν	ψ^1	ψ^2	ν	ψ^1	ψ^2	ν
		France			Germany			Italy		
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	0.390	0	0.739	0.261	0	1	0	0
		(0.264)	(0.264)		(0.163)	(0.163)				
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	0.948	0	0.155	0.845	0	0.219	0.781	0
		(0.187)	(0.187)		(0.241)	(0.241)		(0.218)	(0.218)	
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
		Spain			UK			US		
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	0.156	0
								(0.113)	(0.113)	
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	0.417	0	0.776	0.224	0	0.019	0.981	0
		(0.214)	(0.214)		(0.151)	(0.151)		(0.108)	(0.108)	
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

up for about five quarters for Italy and the UK and then fades away within 3 years, while a longer building up time is found for Spain (10 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 nonnegligible 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.

Transitory dynamics in house prices

As shown in Section II, the permanent-transitory decomposition obtained from the common trends

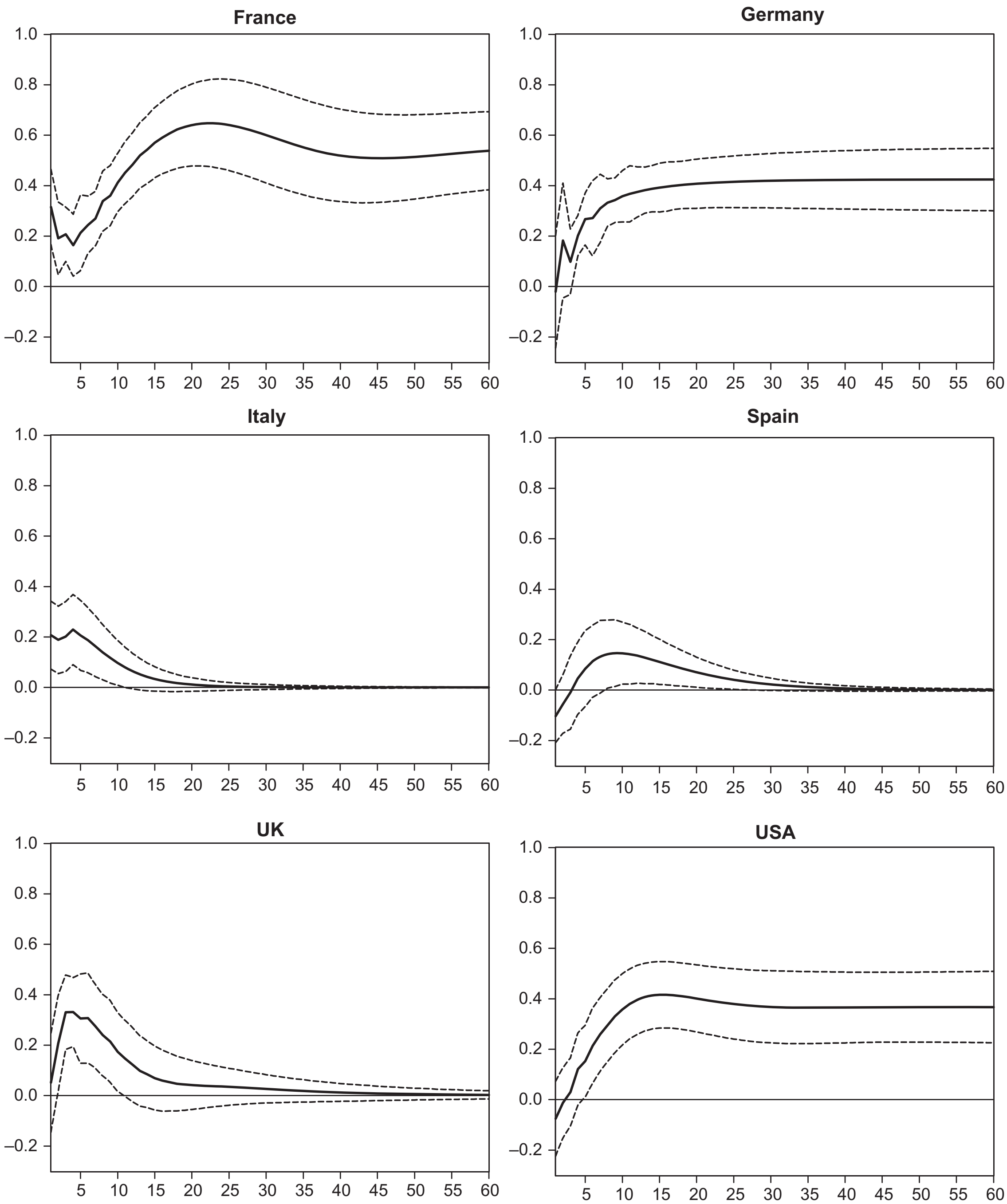

Fig. 4. Consumption response to a permanent shock ψ^2

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

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 Equation 15 the transitory component can be written as

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

where $\Gamma_{2j}^*(L)$ $(j=1,2,3)$ are the elements of the second row of matrix $\mathbf{\Gamma}^*(L)$ in Equation 15. The first two terms on the right-hand side of Equation 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 Fig. 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 *et al*., 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 the error correction properties, is in fact in contrast with the nonstationary growth in house prices, which would be implied by a process of self-fulfilling expectations, spreading, for instance, according to social epidemics (Shiller, 2007).

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 *et al*. (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 *et al*. (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 article, 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

[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 *et al*. (2005) and McCarthy and Peach (2004), who also control for demographic factors.

[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.

recent results of Bai (2003, 2004) and Bai and Ng (2004),[11] principal components analysis has been employed and the results reported in Table 6.

There is a 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 (40–70%), 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 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.

IV. Implications for the Real Effects of the Subprime Crisis

Three main conclusions can be drawn from the empirical analysis carried out so far. First, house prices affect significantly private consumption. While for Italy, Spain and the UK the linkage holds in the short- and medium-term only, for the US, France and Germany long-term effects are detected. Second, house prices tend to be mainly determined by permanent shocks. Only for the US transitory shocks seems to have played a sizeable role during the last expansionary phase. Finally, house price dynamics reflects both global and regional factors, with international synchronization being stronger at the trend, rather than at the cycle, level. Recent results by Beltratti and Morana (2008) also point to the US as a source of global shocks, not only for real activity, but also for inflation and financial variables (interest rates, money supply, real stock and house prices). As the current crisis has shown, indeed global (US) shocks may spill over quickly to other industrialized countries.

In order to better measure the response of consumption to changes in house prices determined by the (permanent) output-neutral shock, already shown in Fig. 4, and then investigate some of the implications of the current crisis for real activity, the impulse response function of c to ψ 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 1% increase in real house prices

[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 heteroscedasticity in both the time and cross section dimension in the idiosyncratic components. In Bai (2004) consistency and asymptotic normality has been derived for the case of $I(1)$ unobserved factors and $I(0)$ idiosyncratic components, also in the presence of heteroscedasticity 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.

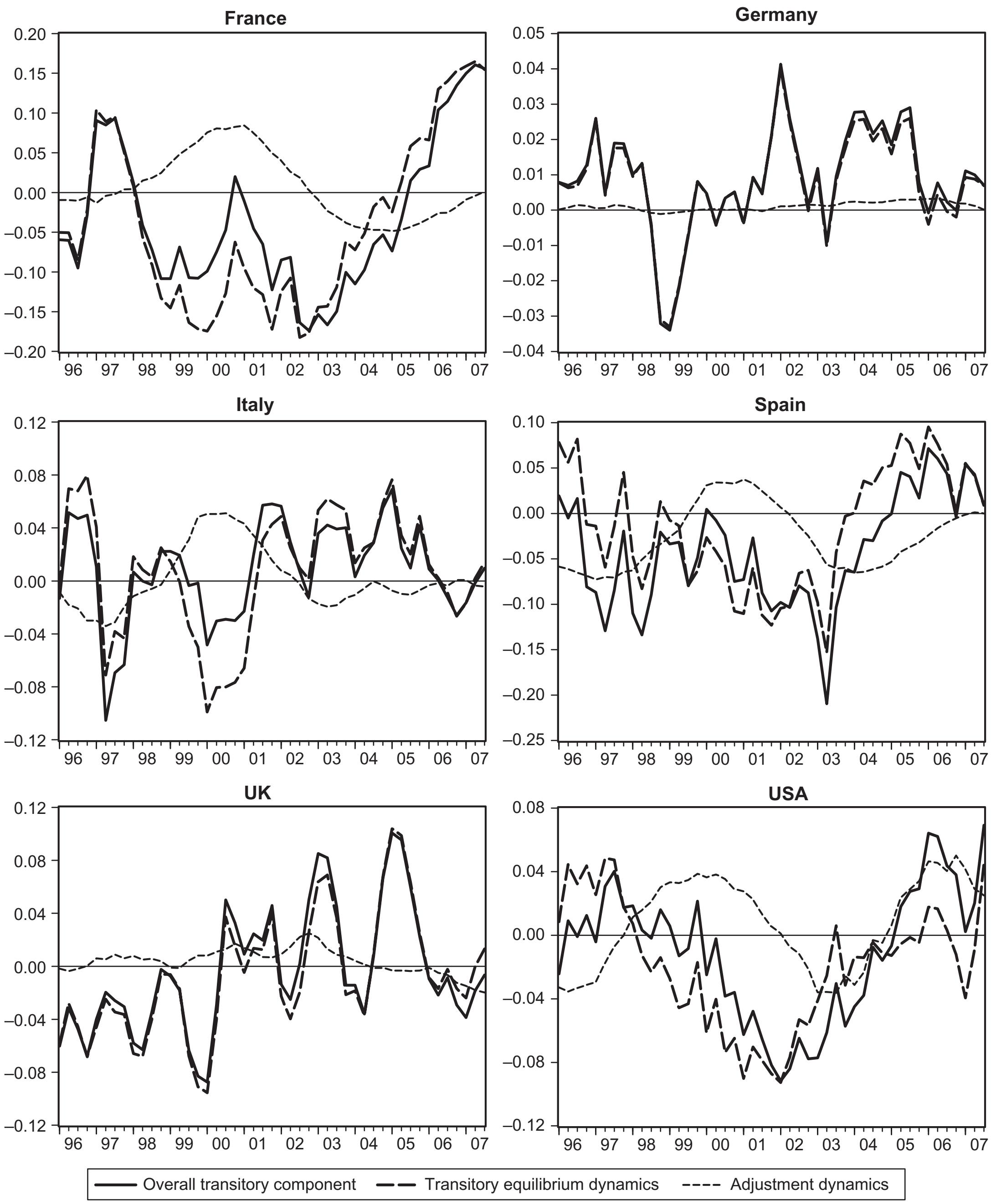

Fig. 5. Analysis of transitory house price component: 1996–2007

due to the output-neutral shock. In Panel A of Table 7, the normalized figures are reported for various horizons of interest, ranging from 1–5 years (for completeness, also the infinite-horizon elasticities, 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 ranging from 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 2-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

Table 6. 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

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

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 1-year horizon very close to that of US consumption expenditure.

The above estimates can be used to get some insights on the potential impact on consumption expenditure of the current pronounced fall in real house prices. According to one well-known indicator, the Case–Shiller index, US real house prices have contracted by about 24% in 2008. Using this figure as an estimate of the expected contraction in house prices also for the European countries and the elasticities reported in Panel A of Table 7, it is possible to calculate the resulting decline in consumption expenditures over the 1-year to 5-year horizons (and the long-run, infinite-horizon impact), as shown in Panel B of the table. In most countries, the peak effect on the level of aggregate consumption expenditure occurs over a medium-term horizon of 2–4 years, reaching −3.4% in France, −6% in Germany, −4.6% in the US and −1% in Spain. In Italy and the UK the peak effect occurs at a 1-year horizon (−1.4% in Italy and −2.9% in the UK).

Table 7. Potential consumption response to real house price shocks

Horizon (quarters)	4	8	12	16	20	∞
Panel A						
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
Panel B						
France	−3.1	−3.4	−3.4	−3.1	−3.1	−3.1
Germany	−4.3	−5.5	−5.8	−6.0	−6.0	−6.2
US	−3.4	−4.6	−4.6	−4.4	−4.1	−3.8
Italy	−1.4	−0.7	−0.2	−0.2	0.0	0.0
Spain	−0.2	−1.0	−0.7	−0.7	−0.5	0.0
UK	−2.9	−1.7	−1.0	−0.5	−0.5	0.0

Notes: Panel A reports the percentage response of consumption expenditure to a 1% change in real house prices due to the output-neutral permanent shock ψ^2. Panel B reports the percentage response of consumption expenditure to a −24% permanent change in real house prices.
* Denotes statistical significance at the 5% level.

Of course, those calculations are subject to many caveats. First of all, they assume that all the observed house price decline is due to permanent disturbances, whereby providing a maximum estimate of the real effect on consumption of house price reductions. Moreover, they neglect the potential effects of the expansionary fiscal and monetary policies currently implemented in all countries under study. Finally, the additional wealth effects stemming from stock price developments are not captured by the reported calculations. Yet, the above estimates may be taken as a rough but useful benchmark to evaluate the expected impact of the financial crisis on consumption spending, and do not appear to be quantitatively unreasonable in the light of the current unwinding of the recession, that featured in 2008 consumption declines by 0.4% in Germany, 1.8% in the US, 2.3% in Spain, 1.4% in Italy, 1% in the UK, and positive growth only in France by a small 0.7%.

V. 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 analysed 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 3–5-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 an evidence of a linkage between house prices and consumption over shorter (around 1–2 years) horizons.

On the basis of the estimated elasticities, a measure of the expected medium- to long-term impact on consumption expenditure of a decline of real house prices of the magnitude experienced by the US in 2008 (−24% according to the Case–Shiller index) is obtained. The results show large effects for the US and Germany (−4.6% and −6%, respectively, at the peak), intermediate effects for France and the UK (−3.4% and −2.9%) and relatively small effects for Italy and Spain (−1.4% and −1%). Though subject to many caveats (most importantly the assumption that all the observed house price decline has a permanent nature), those results may provide a useful (maximum) benchmark measure of the expected real impact on consumption expenditure of the crisis currently hitting housing markets.

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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