Journal of Housing Economics 19 (2010) 38–50

Contents lists available at ScienceDirect

Journal of Housing Economics

journal homepage: www.elsevier.com/locate/jhec

Macroeconomic determinants of international housing markets

Zeno Adams, Roland Füss *

Union Investment Chair of Asset Management, European Business School (EBS), International University Schloss Reichartshausen, D-65375 Oestrich-Winkel, Germany

ARTICLE INFO

Article history:
Received 15 August 2008
Available online 21 October 2009

JEL Classification:
C33
G15

Keywords:
Housing market
Macroeconomy
Panel cointegration
Dynamic OLS
Dynamic adjustment process

ABSTRACT

This paper examines the long-term impact and short-term dynamics of macroeconomic variables on international housing prices. Since adequate housing market data are generally not available and usually of low frequency we apply a panel cointegration analysis consisting of 15 countries over a period of 30 years. Pooling the observations allows us to overcome the data restrictions which researchers face when testing long-term relationships among single real estate time series. This study does not only confirm results from previous studies, but also allows for a comparison of single country estimations in an integrated equilibrium framework. The empirical results indicate house prices to increase in the long-run by 0.6% in response to a 1% increase in economic activity while construction costs and the long-term interest rate show average long-term effects of approximately 0.6% and −0.3%, respectively. Contrary to current literature our estimates suggest only about 16% adjustment per year. Thus the time to full recovery may be much slower than previously stated, so that deviations from the long-term equilibrium result in a dynamic adjustment process that may take up to 14 years.

1. Introduction

In the past, long-run equilibrium models of housing markets and the macroeconomy have been restricted to countries that offer a high availability of long time series of housing market data since cointegration techniques such as the Engle–Granger or the Johansen approach require a sufficiently large time period in order to test for long-run relationships. Accordingly, most studies have focused primarily on the U.S. (see e.g. Case, 2000; Meen, 2002; McCarthy and Peach, 2004), the U.K. (Bowen, 1994; Meen, 2002; Hunt and Badia, 2005) and a few other countries.[1] To overcome this restriction we use macroeconomic and housing market data from 15 OECD countries and apply the panel cointegration approach proposed by Pedroni (1999, 2004). This method makes use not only of the T observations of a time series of a single country but pools the observable data over all N countries so that in effect $N \cdot T$ real observations are available for estimation. This results in a higher robustness of the estimation process since the effects of large sample asymptotics are more likely to apply in this setting. The main advantage, however, lies in the possible finding of an international housing market result which can be obtained by weighting the individual countries' estimations. This result not only allows for predicting house prices on an aggregate international level but also to observe differences among countries' elasticities and thus to draw conclusions on the level of cross-country integration. Finally, recent studies have found that panel data variables are cointegrated even when there was no cointegration between them in individual time series. Hence, this study does not only confirm results from earlier studies but also highlights the differences between countries in an integrated long-run equilibrium framework.

* Corresponding author. Fax: +49 (0) 6723 991 216.
E-mail addresses: zeno.adams@ebs.edu (Z. Adams), roland.fuess@ebs.edu (R. Füss).

[1] For a recent and comprehensive literature overview see Leung (2004).

1051-1377/$ - see front matter © 2009 Elsevier Inc. All rights reserved.
doi:10.1016/j.jhe.2009.10.005

The study is organized as follows: The next section positions our paper within the current stand of existing literature and briefly discusses the choice of variables for estimating house price elasticities. We posit that the slow propagation mechanism of macroeconomic impacts and feedback effects among variables are the key requirements for looking at long-term relationships. Section three presents the panel cointegration technique for non-stationary panel data which we use to test and estimate long-term equilibrium relationships between the housing market and macroeconomic variables. Furthermore, an error correction model is estimated to describe the adjustment process to this long-term equilibrium. Some conclusions are drawn in Section 4.

2. Macroeconomic effects and their propagation mechanism on the housing market

In contrast to other capital market assets, real estate prices do not change immediately after economic news have been released and generally exhibit low price fluctuation. Residential house prices particularly exhibit strong downward price stickiness since homeowners have high reservation prices or simply resist selling their house below a certain price during recessions. Thus, real house prices tend to decrease through inflation rather than through nominal price reductions. Price inertia, however, also influences the behavior of housing prices during economic booms since exuberant expectations of house owners facilitate the formation of housing bubbles.[2] Case (2000) and Catte et al. (2004), among others, have studied the propagation of macroeconomic shocks on U.S. house prices. Macroeconomic shocks such as unexpected changes in the money supply, industrial production, or interest rate changes affect house prices with a lag depending on the speed of the propagation mechanism. The speed of propagation is strongly influenced by the efficiency of the institutional framework such as the land availability, zoning regulations, and the speed of administrative processes. Other variables such as credit supply, transaction costs, and mortgage product innovations also play a major role. For instance, if changes in interest rates propagate quickly into changes of mortgage market interest rates then an increase in the money supply affects housing markets much faster than in a situation where most mortgage rates are fixed and the mortgage market is generally inefficient. The credit supply for housing finance also varies among countries depending on the real estate valuation methods. If the valuation method reacts sensitively to changes in real estate prices and if the loan-to-value (LTV) ratio[3] is high then rising house prices increase the credit supply more strongly and, vice versa, decreasing house prices lead to a shortage in the credit supply. A higher credit supply, in turn, increases the importance of interest rate changes as more firms and households rely on debt financing. Lower transaction costs lead to more transactions and thus to a faster response of house prices in face of a macroeconomic shock. On the other hand, if the response of house prices is very low strong price reactions are more likely.

House prices may also exhibit a feedback reaction to the macroeconomy. Rising house prices make homeowners feel richer[4] because the value of their houses and thus the size of their collateral they can borrow on increases. For liquidity-constrained households an increase in house prices may be their only opportunity to borrow at all. This wealth effect then increases consumption. A decline in house prices leads to a negative effect on consumption since decreasing house prices lead to more mortgage defaults and thus reduce the supply of bank credit as banks lose part of their bank capital (Parker, 2000). Here, the mortgage market also plays an important role for the propagation of real house prices to the macroeconomy. Higher mortgage debt means higher leverage, through which changes in the interest rate can affect consumer spending. As Case (2000) finds, the effect of house prices on consumption is especially strong in the U.S., where two thirds of all occupants are also owner-occupants so that the wealth effect has a strong impact on consumer spending. Case et al. (2005) show that changes in real house prices can even impact consumption more strongly than changes in stock market prices, which might be due to the fact that house ownership is more evenly distributed across households than stock market wealth. In contrast, stock market wealth is mainly held by rich households. Since the propensity to consume declines with increasing wealth an increase in house prices should therefore have a stronger effect on consumption than an increase in stock prices.[5]

Case (2000), Sutton (2002), Tsatsaronis and Zhu (2004), and Terrones and Otrok (2004) have studied global macroeconomic effects on real estate prices. Real estate markets appear to be highly correlated internationally although, being bound to a specific location, they are not substitutes. However, fundamentals like GDP which drive real estate markets are internationally correlated.[6] The strength of those global factors depends on the openness of the country. GDP correlations were found to range on average from 0.33 to 0.44 (Case, 2000).[7]

The slow propagation of macroeconomic effects on the one hand and endogenous feedback effects on the other

[2] Furthermore, price information in the real estate market is often limited and inaccurate as real estate is sold only infrequently and information about prices is often specific to the respective local market. For this reason, price forecasts are usually simple extrapolations from the past. This leads to endogeneous dynamics of real estate prices and thus facilitates the formation of housing bubbles. The contribution of those endogeneous dynamics can be substantial and varies between 70% and 40% depending on the country under consideration (Zuh, 2003).

[3] Loan-to-value (LTV) is defined as ratio of the value of the loan to the value of the collateral.

[4] For instance two thirds of the U.S. population are homeowners. The increase in house prices is of course disadvantageous for the other one third of the population who rent houses as rents increase as well. Feedback effects have also been found in various countries, e.g., Campbell and Cocco (2007) for the UK, Berg and Bergström (1995) for Sweden, and Koskela et al. (1992) for Finland.

[5] Some economists, however, do not believe in the existence of such wealth effects. See, e.g., Glaeser (2000).

[6] For example, the real estate crash in office prices of the early 1990s and the residential house price crash in 2008 were felt by nearly all countries around the world.

[7] Other non-macroeconomic effects have been studied as well. See, e.g., Parker (2000) for the effect of population growth on house prices and Cocco (2005) for a discussion of real estate in the portfolio context.

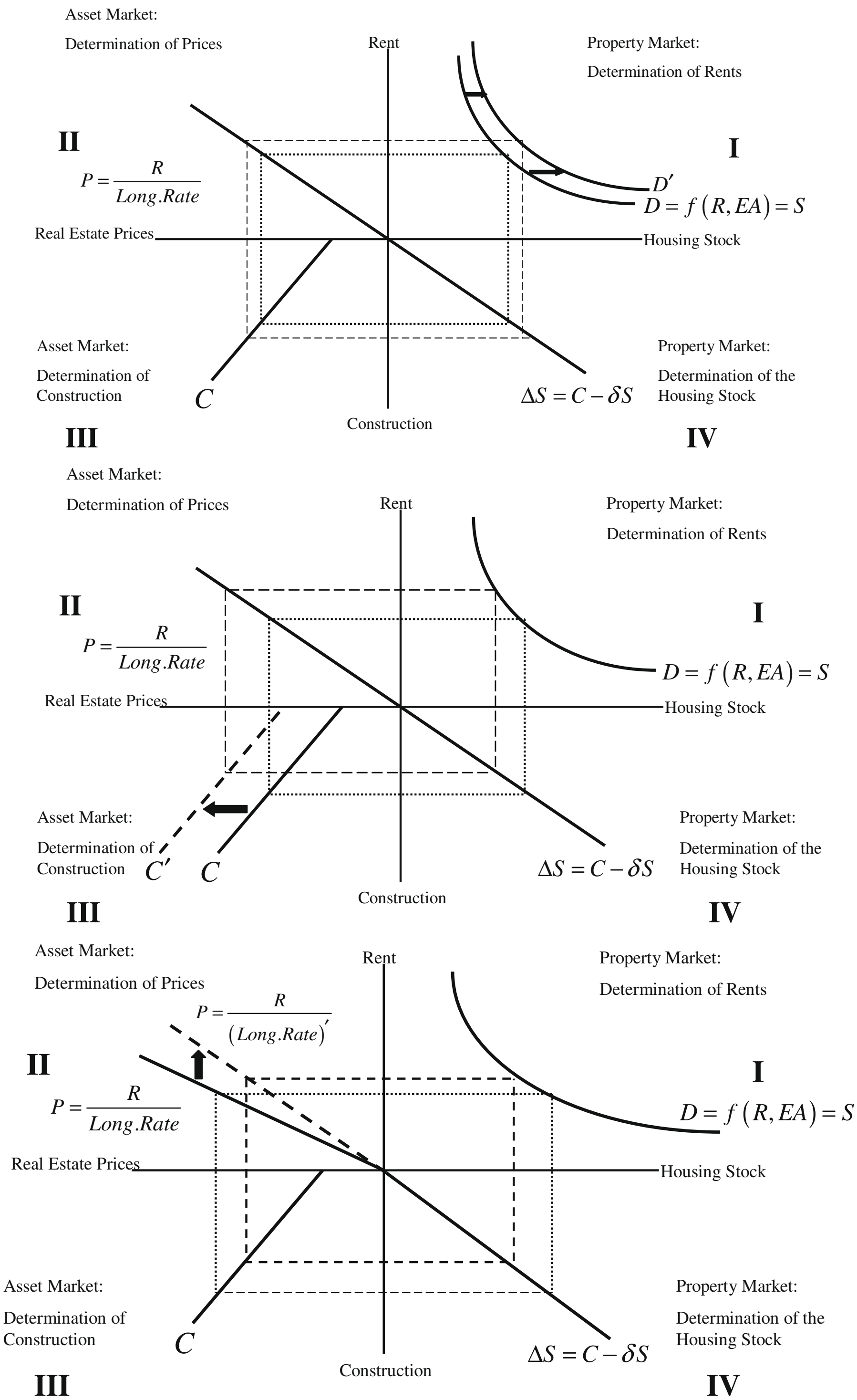


Fig. 1. The impact of macroeconomic variables on real house prices. The upper panel shows the effects of an increase in economic activity. The middle panel shows the effects of an increase in long-term interest rates. The lower panel shows the effects of an increase in general construction costs.

may bias estimates of contemporaneous elasticities. What is asked for are methods with long-term perspectives that estimate elasticities when all endogenous feedback effects have been worked out.

The choice of which variables to use for estimating the driving factors of house prices largely depends on the underlying models. These can be roughly segmented into econometric models, affordability indicators, and asset-pricing approaches (Girouard et al., 2006). Since this study does not primarily aim at identifying overvaluations in house prices affordability indicators such as the price to rent ratio or various indebtedness measures were not used. The benefit of asset-price approaches is to model an equilibrium condition that arises from the simple arbitrage relationship that housing rents should equal the user cost of housing. Deviations from this equilibrium condition are interpreted as error correction terms that lead to adjustments in the following periods. In this study we follow the current stand of econometric models which often choose their variables in an ad hoc but plausible manner. Here, the adjustment dynamics that eventually restore equilibrium are modeled econometrically. Our choice of variables for the estimation process is based on the static equilibrium model of DiPasquale and Wheaton (1996), hereafter DW. Central to this model, which is displayed in Fig. 1, is the distinction between the asset and the property market as well as their interaction.

This model leads us to the choice of the following three variables

1. Economic activity: The upper panel in Fig. 1 shows the effects of an increase in economic activity, EA. An increase in EA through, e.g., an increase in employment, or real industrial production increases the demand for space, *D*, and shifts the demand curve in the first quadrant to the right. Since the housing stock cannot change in the short-run, rents, *R*, increase leading to higher house prices, *P*, in the asset market: $\underbrace{\frac{\partial P}{\partial \mathrm{EA}}}_{+} = \underbrace{\frac{\partial R}{\partial D}}_{+} \cdot \underbrace{\frac{\partial D}{\partial \mathrm{EA}}}_{+}$. A widely used indicator for economic activity is disposable income. However, this variable is a measure of average income whereas house buyers and sellers typically have incomes that are higher than the population mean. Therefore, we generate the variable "Economic Activity" by calculating the first principal component of the matrix consisting of real money supply, real consumption, real industrial production, real GDP, and employment.

2. Long-term interest rate: An increase in the long-term interest rate does not change the demand for housing space directly but changes the demand to own houses. A higher long-term interest rate, *i*, increases the return of other fixed-income assets such as bonds relative to the return of real estate, thus shifting the demand from real estate into other assets. A higher long-term interest rate is furthermore reflected in higher mortgage rates, which reduces demand and further decreases house prices.[8] This change in demand is shown in the middle panel of Fig. 1 and results in an increase in the slope of the capitalization rate which is the ratio of rents to house prices.[9] The higher capitalization rate is reflected in lower real estate prices which in turn decreases construction and thus translates into a lower housing stock. The lower housing stock increases rents so that the new box is higher and more symmetric than the previous one. We would therefore expect a negative effect on house prices in the size of $\frac{\partial P}{\partial i} = -\frac{R}{i^2}$. Note that it is sufficient to include the long-term interest rate in nominal terms. As inflation reduces nominal returns from an investment in real estate and other fixed-income assets in the same way, the relative attractiveness between both asset types remains unchanged. Furthermore, we depart from the usual setup in asset-pricing types of house price models by neglecting short-term interest rates. This is because, theoretically, the effects of short-term interest rates are ambiguous: An increase in the short-term real interest rate reduces house prices through lower demand for housing due to higher mortgages rates for ARMs. On the other hand, however, interest rates could also increase house prices because higher financing costs reduce housing supply. In contrast, the long-term interest rate has no such ambiguous effects. Indeed, the negative effects on house prices induced by capital switching are most likely intensified by higher interest rates for fixed rate mortgages.[10]

3. Construction costs: The third effect likely to impact the supply schedule of new construction is a change in the construction costs. Higher costs of construction, such as an increase in the price of construction materials or higher labor costs increase the financing costs of construction. This effect is shown in the bottom panel of Fig. 1, where the construction line shifts to the left. The higher construction costs lead to a decrease in construction, *C*, and thus to a lower level of the housing stock, *S*.[11] The lower housing stock also means less housing space which increases rents. Higher rents then generate higher house prices in the asset market: $\underbrace{\frac{\partial P}{\partial C}}_{+} = \underbrace{\frac{\partial R}{\partial S}}_{-} \cdot \underbrace{\frac{\partial S}{\partial C}}_{-}$. As evident, the location of the new box is higher and more to the left than the previous box. Rents and house prices are higher but construction and the housing stock are lower than they would be without the increase in construction costs. The exact position of the box depends on the elasticities of the individual curves.

Before leaving the issue, note that we considered a demographic variable as well but ultimately decided against it. Since the frequently cited paper by Mankiw and Weil (1989) most studies have found insignificant or negative effects of population growth on house prices (e.g., Berg, 1996; Hort, 1998 for Sweden, and Engelhardt and Poterba, 1991 for Canada; Poterba, 1991).

Within the DW framework we thus implement the following long-run model of supply and demand. The demand function is given as

[8] The long-term interest rate affects mainly currently closed fixed rate mortgage contracts while adjustable rate mortgages (ARMs) are primarily affected by short-term interest rates. Despite a recent trend towards ARMs, fixed rate mortgages are still the main borrowing vehicle in most countries (for a country overview, see Girouard et al., 2006).

[9] The capitalization rate is defined as $cap = (NOI - debt)/sales\ prices$, where NOI is net operating income (gross income minus operating expenses) and *debt* is debt service payments. This earnings-price ratio is often considered relative to that of other investments.

[10] We also estimated a model that included both, short-term and long-term interest rates. Despite of some evidence for short-term interest rates affecting house prices in a positive way and long-term interest rates having a negative effect on house prices the estimated effects suffered from the strong multicollinearity between the two rates. In this paper, we therefore focus on long-term interest rates.

[11] Thus, in our case it is not necessary to take the housing stock explicitly into account.

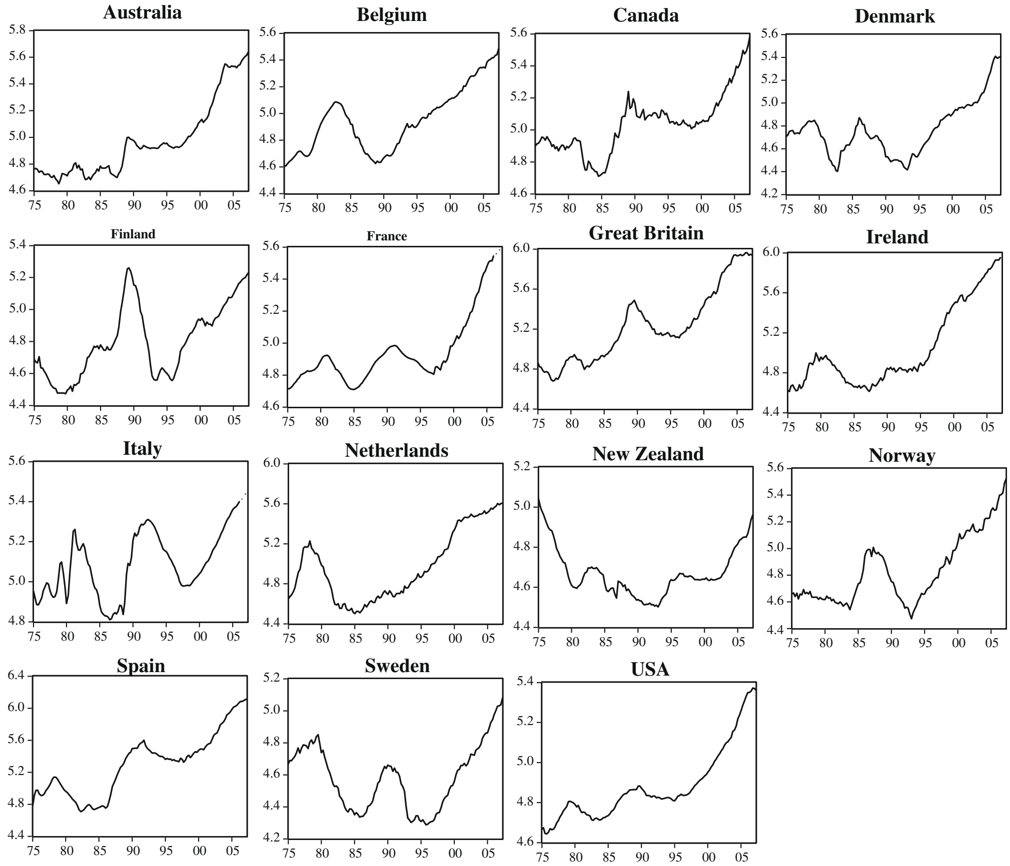


Fig. 2. Log real house prices 1975q1 – 2007Q2. The graphs show the quarterly log real house price series, i.e., nominal house prices deflated by national consumer price indices.

$$D_t = \alpha + \beta' x_t^D + \delta' z_t^D + \varepsilon_t, \tag{1}$$

where x_t^D is a vector of macroeconomic variables affecting demand. Vector z_t^D captures country-specific factors affecting housing demand at the micro level, such as distance to the CBD, social environment, mortgage market characteristics, and tax regulations. Since we focus in this study on the macroeconomic impact on housing markets, the vector z_t^D is incorporated into the error term and Eq. (1) can be written as[12]

$$D_t = \alpha - \beta_1 hp_t + \beta_2 \mathrm{EA}_t - \beta_3 long_t + \tilde{\varepsilon}_t. \tag{2}$$

In Eq. (2), higher house prices hp_t decrease demand for home ownership, while higher economic activity EA_t has a positive effect on demand.[13] Higher long-term interest rates make fixed-income assets more attractive relative to housing investments and thus lead to capital switching which lowers demand for home ownership. In addition, an increase in the long-term interest rate reduces demand through higher mortgage rates. In a similar fashion housing supply is given by

$$S_t = \eta + \gamma' x_t^S + \lambda' z_t^S + \nu_t. \tag{3}$$

And the supply equation can be expressed in more detail as

$$S_t = \eta + \gamma_1 hp_t - \gamma_2 \mathrm{constr}_t + \tilde{\nu}_t, \tag{4}$$

which incorporates micro factors such as the availability of land or governmental building provisions into the error term, $\tilde{\nu}_t$. In Eq. (4) higher house prices act as an incentive for investors to increase the supply of houses while higher construction costs constr_t have a negative impact on housing supply.[14] The equilibrium relationship is established by equating supply and demand. Solving for house prices and considering the panel structure, we have

$$hp_{it} = \alpha_i^* + \beta_{2i}^* \mathrm{EA}_{it} + \gamma_{2i}^* \mathrm{constr}_{it} - \beta_{3i}^* long_{it} + \varepsilon_{it}^*, \tag{5}$$

with

$\alpha_i^* = \frac{\alpha_i - \eta_i}{\gamma_{1i} + \beta_{1i}}$, $\gamma_{2i}^* = \frac{\gamma_{2i}}{\gamma_{1i} + \beta_{1i}} = \frac{\partial P_i}{\partial C_i}$, $\beta_{2i}^* = \frac{\beta_{2i}}{\gamma_{1i} + \beta_{1i}} = \frac{\partial P_i}{\partial EA_i}$, $\beta_{3i}^* = \frac{\beta_{3i}}{\gamma_{1i} + \beta_{1i}} = \frac{\partial P_i}{\partial i_i}$, and $\varepsilon_{it}^* = \tilde{\varepsilon}_{it} - \tilde{\nu}_{it}$.

[12] Although we consider the micro factors to be relevant they cannot be taken into account in a panel of 15 aggregated country indices.

[13] During the subprime bubble, rising house prices may have been regarded as a factor that *increased* demand for homeownership. Our dataset includes the beginning of the bubble period but, as we show later, the effects are not fundamentally different when working with a shorter time period that excludes the bubble.

[14] The construction cost index is composed of a wide range of cost items, including costs for and transportation of building materials, labor costs, and interest on loans.

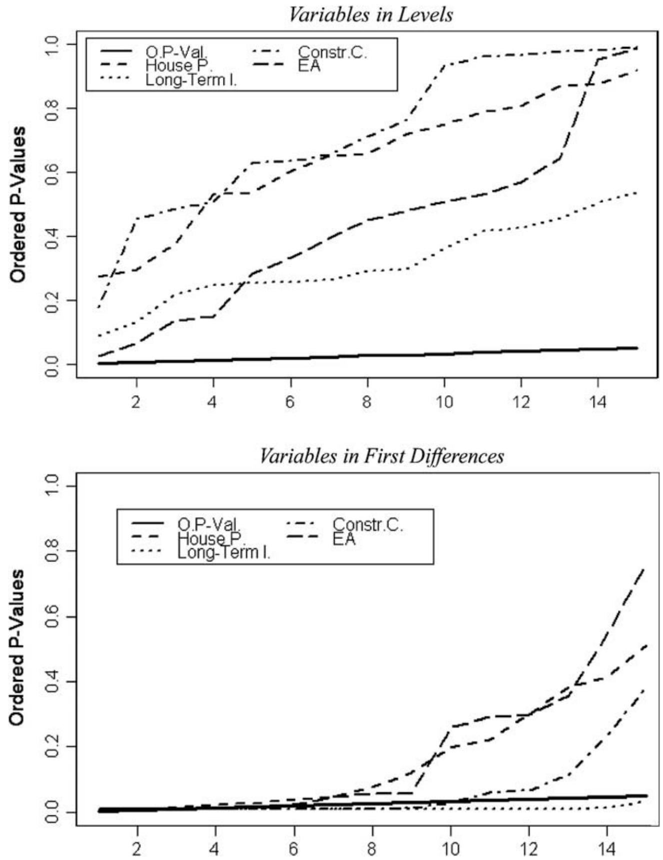


Fig. 3. Hanck panel unit root test. Calculations are based on quarterly data from 1975Q1 to 2007Q2. The null hypothesis of a unit root can be rejected if at least one ordered p-value is below the critical value of $i{\cdot}0.05/j$ for $i = 1, 2,\ldots,15$, and $j = 15$ (thick solid line).

In the context of the theoretical model discussed above, we expect a positive sign for β_2^* and γ_2^*, and a negative sign for β_3^*.

3. Long-term equilibrium relationships between housing markets and the macroeconomy

Cointegration analysis for non-stationary panel data is conducted in three steps: First, we test the variables for stationarity using panel unit root tests. Next, we apply panel cointegration tests to detect the long-term equilibrium relationships, and finally we estimate the short-term dynamics.[15] For the estimated models below we use data from 1975Q1 to 2007Q2 for the following 15 countries: Australia, Belgium, Canada, Denmark, Finland, France, Great Britain, Ireland, Italy, the Netherlands, New Zealand, Norway, Spain, Sweden, and the U.S.[16] We use the *BIS house price* data which are derived from various national sources and thus do not provide the same quality for all countries. Fig. 2 shows the development of the log real house prices since 1975.[17]

The macroeconomic factors that enter our models as independent variables are *economic activity*, *long-term interest rates*, and *construction costs*. The variable *economic activity* is created by the first principal component of the matrix consisting of real money supply, real consumption, real industrial production, real GDP, and employment.[18]

3.1. Panel unit root tests

Early work on non-stationary panel data include Quah (1994) or Levin and Lin (LL) (1993) who study unit root tests under the null hypothesis of non-stationarity assuming the parameters of the lagged endogenous variable ρ_i to be homogeneous for all individuals,

[15] Cointegration methodology for testing long-term equilibrium relationships between single time series have been developed by Engle and Granger (1987) for the univariate case and by Johansen and Juselius (1990) for the multivariate case.

[16] Germany and Japan were excluded from the data set due to inaccurate house price data.

[17] Although we are aware of the potential impact of the housing bubble on our analysis, we decided to include the recent years in order to include recent relevant information and to increase the total number of observations.

[18] The data for the macroeconomic variables come from Datastream. When possible we used the same data source for all countries so that most macroeconomic data is provided by the OECD. Long-term interest rates are 10 year government bonds.

$$\begin{aligned}&\Delta y_{it} = \alpha_i + \delta_i t + \rho_i \Delta y_{it-1} + e_{it}\\&\qquad \text{for } t = 1,\dots,T \text{ and } i = 1,\dots,N,\\&H0: \rho_1 = \rho_2 = \ldots = \rho_N = \rho = 0\\&\qquad \text{against the alternative hypothesis}\\&H1: \rho_1 = \rho_2 = \ldots = \rho_N = \rho < 0.\end{aligned} \tag{6}$$

Although the LL test allows for heterogeneity in the variance and serial correlation structure of the error terms, the restriction of homogeneous slope parameters is clearly too strong and the alternative hypothesis is thus of no practical interest. Im, Pesaran and Shin (IPS) (1997) and Maddala and Wu (1999) propose unit root tests that also allow for heterogeneous autoregressive roots. The IPS test is a generalization of the LL test that combines the test statistics of the individual unit root tests for each cross-section unit. The IPS test has the advantage of heterogeneous slope parameters so that the alternative hypothesis accordingly becomes

$$\begin{aligned}&H1: \rho_i < 0 \quad \text{for } i = 1,2,\dots,N_1 \text{ and}\\&\rho_i = 0 \text{ for } N_1+1, N_2+2,\dots,N,\end{aligned}$$

so that slope parameters ρ_i are allowed to differ across group members and not all N members need to be cointegrated.

So called 'second generation panel unit root tests' (Phillips and Sul, 2003; Moon and Perron, 2004; Breitung and Das, 2005; Hanck, 2008) are also reliable in the presence of cross-sectional dependence. In this paper we apply the Hanck (2008) test, which has been shown to exhibit good size and power properties and is easy to implement as the p-values of the single time series unit root tests are sufficient as inputs. Denoting $p_{(1)}, p_{(2)}, \dots, p_{(n)}$ as the ordered p-values $p_{(1)} \leqslant p_{(2)} \leqslant \ldots \leqslant p_{(n)}$, the null hypothesis is rejected at significance level α if and only if

$$\exists j \epsilon \mathbb{N}_n : p_j \leqslant j \cdot \alpha / n. \tag{7}$$

Here, $n = 15$ and the critical value $\alpha = 0.05$.

As shown in Fig. 3, employing p-values from the single time series Phillips–Perron test the null hypothesis could not be rejected for all four variables although a few countries show a stationary process for some variables.[19] Taking first differences reveals stationarity in the panel setting so that all variables can be considered as $I(1)$ and are thus suitable for the cointegration test procedure.

3.2. Panel cointegration test

If the macroeconomic and housing market variables are $I(1)$ but a linear relationship between those variables is $I(0)$, the variables are cointegrated. In order to test for cointegration, we apply a cointegration test for heterogeneous panels with multiple regressors which was developed by Pedroni (2000). This test has the null hypothesis of no cointegration and also allows for unbalanced panels. In this test the regression residuals are computed from the regression

$$\begin{aligned}&y_{it} = \alpha_i + \delta_i t + \gamma_{1i} x_{1it} + \gamma_{2i} x_{2it} + \ldots + \gamma_{Mi} x_{Mit} + e_{it}\\&\quad \text{for } t = 1,\dots,T; i = 1,\dots,N; m = 1,\dots,M\end{aligned} \tag{8}$$

with individual fixed effects α_i and individual time trends $\delta_i t$, although such individual time trends are often omitted. In some cases, common time dummies can also be included. In this equation, m regressors $x_{mi,t}$ are allowed and the slope coefficients ρ_{mi} and thus the cointegration vectors are heterogeneous for all i. The residuals $\hat{e}_{i,t}$ from Eq. (8) are then tested for unit roots,

$$\hat{e}_{it} = \rho_i \hat{e}_{it-1} + \varepsilon_{it}. \tag{9}$$

Including fixed effects and time trends changes the asymptotic distribution and increases the critical values of the unit root statistic. This is because in the presence of a unit root, the sample average of a variable with a stochastic trend $\bar{y}_i = \frac{1}{T}\sum_{t=1}^{T} y_{i,t}$ does not converge to the population mean with increasing T.[20] Pedroni (1999) proposed seven test statistics. For our purpose we decided to use the multivariate extension of the Phillips–Perron Roh (PPr) statistic which is a nonparametric test and does not require the slope coefficients of the regression residuals to be homogeneous for all i in the case of the alternative hypothesis of cointegration. Pedroni (2000, 2004) shows that under general requirements the test statistics follow a normal distribution as T and N grow large. A test statistic of 2.844 for the PPr test rejects the null hypothesis of no cointegration. Thus house prices are cointegrated with economic activity, long-term interest rates, and construction costs.

3.3. Cointegration-vector estimates

The widely used single-equation cointegration-vector estimator proposed by Engle and Granger (1987) is known to have the nice property of super consistency which established the belief that this technique has a certain level of robustness. Super consistency means that the estimated coefficients converge to their true value at rate T instead of $\sqrt{T}$ as the sample size T grows large but says nothing about the statistical properties for a given sample size. Inder (1993) and Stock and Watson (1993) show that these can be quite poor for the typical number of observations available to researchers with macroeconomic time series data.

The Panel Dynamic Ordinary Least Squares (DOLS) estimator introduced by Saikkonen (1991), Phillips and Moon (1999), and Pedroni (2000) augments the conventional OLS estimator by taking serial correlation and endogeneity of the regressors into account. In a series of Monte-Carlo simulations Kao and Chiang (2000) and Mark and Sul (2003) test the small sample performance of the panel DOLS estimator and show that it generally outperforms single-equation estimation techniques.

[19] All variables except for the long-term interest rate have been deflated with the consumer price index, i.e., $x_{\text{real}} = (x_{\text{nominal}}/\text{CPI}) \times 100$. When necessary, the series have been seasonally adjusted using the Census X12 procedure.

[20] In fact, it can be shown that the sample mean diverges at the rate $\sqrt{T}$.

As outlined in Eq. (8) interest lies in the coefficient vector estimate γ' of

$$y_{it} = \alpha_i + \gamma' x_{it} + u^*_{it}, \tag{10}$$

with the regressors x_{it} being integrated of order 1: $x_{it} = x_{it-1} + \upsilon_{it}$.[21] Correlation of the additional error component υ_{it} with u^*_{it} is a potential source of bias and an effective protection from this bias is to explicitly control for this relationship by regressing u^*_{it} on p leads and lags of υ_{it},

$$u^*_{it} = \sum_{s=-p}^{+p} \delta'_{is}\upsilon_{it-s} + u_{it} = \sum_{s=-p}^{+p} \delta'_{is}\Delta x_{it-s} + u_{it} = \delta'_{is} z_{it} + u_{it}, \tag{11}$$

where the second equality follows from $x_{it} = x_{it-1} + \upsilon_{it}$ and the third equality is simply a vectorized notation to conserve space.[22] Inserting this expression into (10) yields the endogeneity and serial correlation-adjusted regression

$$y_{it} = \alpha_i + \gamma' x_{it} + \delta'_i z_{it} + u_{it}, \tag{12}$$

from which the coefficient vector $\beta_{\mathrm{DOLS}} = (\gamma',\ \delta'_1, \ldots, \delta'_N)'$ can be obtained.[23]

The interpretation of the DOLS estimator is now similar to a conventional panel OLS estimator except for one important aspect: A fixed effects estimator would show the response of house prices at time *t* or generally at time *t–p*. The DOLS estimator shows the *long-run* effects, which capture the accumulation of effects over time, as well as the stickiness of house prices. Thus, if all variables are in logs, the elements of the coefficient vector $\gamma' = (\gamma'_{1i}, \gamma'_{2i}, \gamma'_{3i})$ show the average long-term percentage change of the regress and for a one percentage change in the regressor.

Table 1 shows the estimation results for each individual country as well as for the whole panel. In the latter case the coefficients are obtained by averaging over the individual country coefficients. Thus, a 1% increase in economic activity leads on average to a 0.34% increase in house prices over the long run while higher interest rates show a negative impact of 0.4%. A 1% increase in construction costs on the other hand increase house prices by 1.3%.

The results strongly confirm the theoretical implications of the DW model. However, individual coefficients vary widely among countries. For instance, the coefficient for economic activity varies between −0.95 for Canada and +1.06 in the case of the Netherlands.[24] At this stage it is therefore not possible to interpret the panel group coefficients as an international housing market result.

The panel structure offers the opportunity to identify country groups with similar elasticities. Excluding countries with imprecise or wrongly signed coefficients furthermore provides an increase in robustness. The upper panel of Fig. 4 shows the results of a cluster analysis on the country coefficients. From the dendrogram two main groups and one outlier can be identified: Group 1 consisting of the five countries Spain, Canada, Italy, Denmark, and the U.K., and group 2 which includes New Zealand, Ireland, Belgium, Sweden, the Netherlands, the U.S., Australia, Norway, and Finland. France with its negatively estimated effect of economic activity and large coefficient on construction costs is regarded as an outlier. It is interesting to see that group 2 contains mainly small neighboring European countries. On the other hand the U.S., Australia, and New Zealand do not match in this respect and the countries in group 1 are mainly spatially separated. Thus, there seems to be only weak evidence for similar house price reactions among countries with close geographic proximity. From the lower panel of Fig. 2 one can clearly see the strong heterogeneity among the member countries of group 1. Most countries in this group show the wrong sign on the interest rate as well as large elasticities that are not in line with previous research. For example, for Denmark, Wagner (2005) estimates the effects of disposable income and the housing stock to be 2.9 and −2.9, respectively, while our estimates suggest values of around 11 and −1.2. On the other hand, group 2 countries show an average long-run elasticity of 0.61. This value, along with the coefficients for the long-term interest rate (−0.27) and for construction costs (0.64), is generally in line with most of the previous studies, e.g., Mankiw and Weil (1989), Heiborn (1994), Eitrheim (1995), Kosonen (1997), Terrones and Otrok (2004), Annett (2005), which find elasticities of income below unity. In contrast to group 1, the country coefficients in group 2 are furthermore very homogeneous so that these elasticities can be considered as indicative of how international housing markets as a whole may respond to global macroeconomic changes.

3.4. Error correction model

After having analyzed the long-run impact of the macroeconomic variables on the housing market, a natural question would be to ask how long it would take for the housing market to return to equilibrium after an exogeneous shock to the economy. To answer this question, we estimate an error correction model (ecm) which models the adjustment process to equilibrium. Deviations from the equilibrium are expressed as

$$\mathrm{ecm}_{it} = hp_{it} - \beta_1 ea_{it} - \beta_2 constr_{it} - \beta_3 long_{it}, \tag{13}$$

where the β_i are the DOLS parameter estimates. If the variables are in equilibrium, the error correction term ecm_{it} is zero whereas deviations from equilibrium are reflected in a non-zero error term. For example, if house prices hp_{it} are too high relative to their equilibrium value the error term is positive. Under this situation hp_{it} will decrease in the fol-

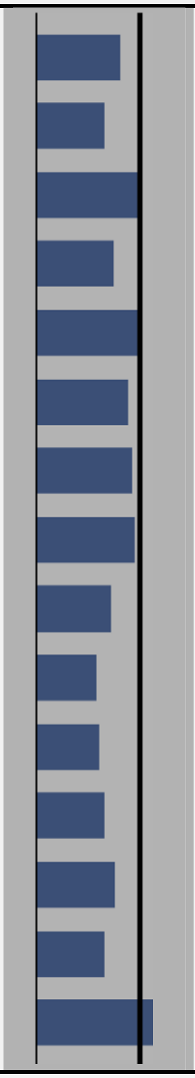
[21] Here, y_{it} is the house price of country *i* at time *t* and x_{it} is a *3x1* vector of economic activity, construction costs, and the long-term interest rate of country *i* at time *t*, respectively.

[22] p is usually chosen to be around 2 for yearly data, i.e. for our calculations using quarterly data p is 8.

[23] Interested readers are referred to Pedroni (2000), Kao and Chiang (2000), and Mark and Sul (2003) for a more detailed mathematical derivation.

[24] The horizontal bars in Table 1 show the amount of variation in the matrix of real money supply, real consumption, real industrial production, real GDP, and employment that can be explained by the variable "Economic Activity". Since the proportions are less than 1, the interpretations of the effects from this variable are not straightforward. Replacing this variable by real GDP for instance turns the coefficients for France, Spain, and Canada to 1.88 (7.75), −0.07 (−0.38) and 2.14 (4.43), respectively, but renders those of Sweden and New Zealand negative.

Table 1
DOLS estimates. Calculations are based on quarterly data from 1975Q1 to 2007Q2. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively; *t*-statistics in brackets. The panel group DOLS results are computed as the averages of the individual country estimates. The lead and lag value *p* is 8 (see Eq. (11)). Estimates are without common time dummies, as house prices are likely to be cointegrated among each other (see, e.g., Terrones and Otrok (2004), who find house prices to be highly synchronized among countries). In this case, time dummies would destroy the cointegration relationship (see Pedroni, 1999). The horizontal bars show the amount of variation in the matrix of real money supply, real consumption, real industrial production, real GDP, and employment that can be explained by the variable "Economic Activity".

Country	Proportions (0.0 – 0.6)	log real economic activity	log (1+long-term interest rate)	log construction costs
Australia		0.30*** [4.40]	-0.01 [-0.19]	0.58*** [4.30]
Belgium		0.06** [2.21]	-0.49*** [-8.69]	0.26*** [3.25]
Canada		-0.95*** [-8.79]	-1.16*** [-5.49]	0.75*** [3.43]
Denmark		0.52*** [5.59]	-0.18*** [-3.77]	2.46*** [9.00]
Finland		0.78*** [5.89]	0.02 [0.16]	0.93*** [10.92]
France		-0.89*** [-14.12]	-0.82*** [-35.42]	4.99*** [17.48]
Great Britain		0.29 [0.92]	-0.17 [-0.80]	1.86*** [5.84]
Ireland		0.68*** [13.82]	-0.42*** [-11.49]	0.09 [1.48]
Italy		0.79*** [19.99]	-0.11*** [-4.41]	2.2*** [40.29]
Netherlands		1.06*** [3.79]	-0.71*** [-4.58]	0.77*** [3.83]
New Zealand		0.56*** [8.56]	-0.27*** [-3.83]	0.28 [1.03]
Norway		0.87*** [13.28]	0.27*** [3.65]	1.17*** [15.42]
Spain		-0.26*** [-2.69]	-1.16*** [-15.42]	1.46*** [7.67]
Sweden		0.99*** [6.90]	-0.45*** [-15.36]	0.59* [1.74]
USA		0.26** [2.09]	-0.33*** [-2.78]	1.11*** [6.27]
Panel Group DOLS Results				
Coefficient		0.34*** [15.96]	-0.40*** [-28.00]	1.30*** [34.07]

lowing periods until the equilibrium is reached. This adjustment process can be estimated by including the lagged error term in a panel regression. In contrast to the long-run cointegration model which has to be estimated with non-stationary variables the short-run error correction model is a conventional panel regression and thus needs to be estimated using stationary variables. All variables except the error term are included as *d*logs,

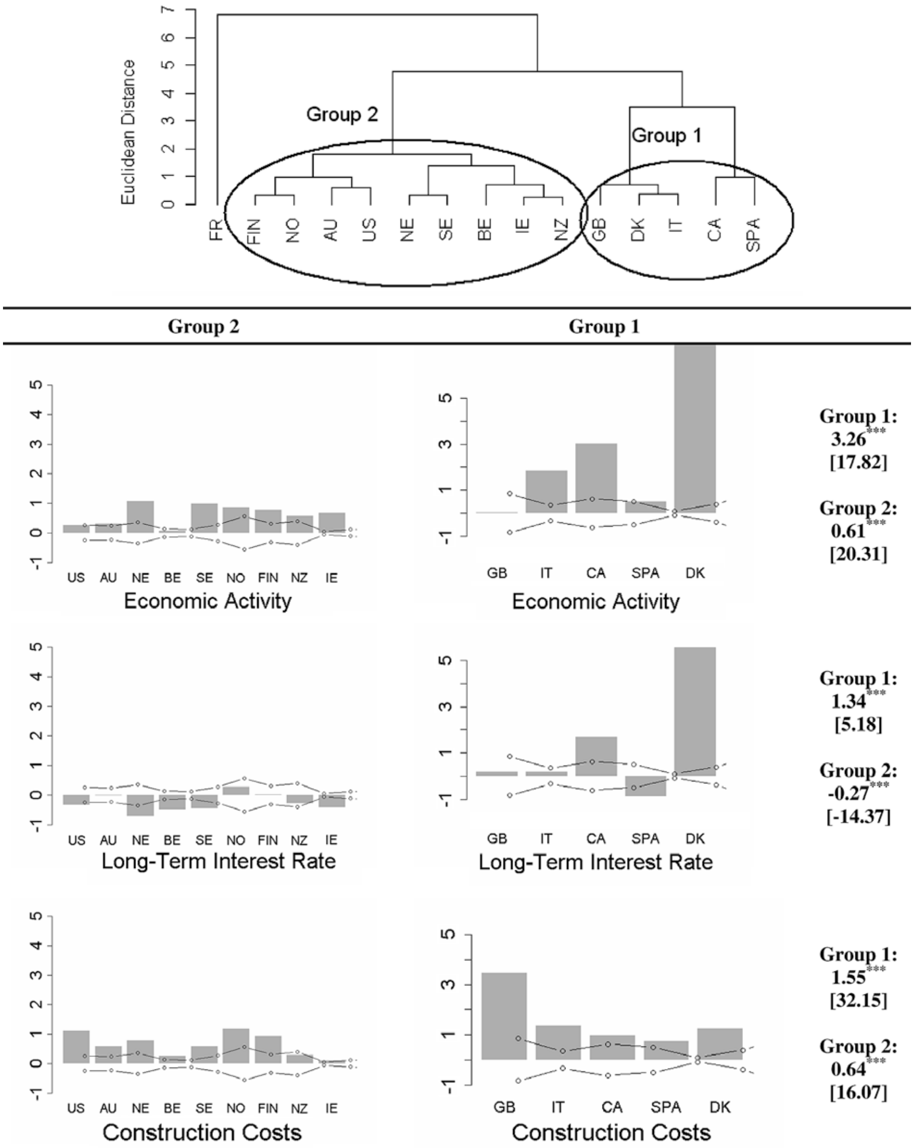


Fig. 4. Similarities among housing market elasticities.

$$\begin{aligned}\Delta\ \log\ hp_{it} = \hat{\alpha}_0 &+ \sum_{j=1}^{K}\hat{\alpha}_{1j}\Delta\ \log\ hp_{it-j} \\ &+ \sum_{n=0}^{N}\hat{\alpha}_{2n}\Delta\ \log\ ea_{it-n} \\ &+ \sum_{s=0}^{S}\hat{\alpha}_{3s}\Delta\ \log\ \mathrm{constr}_{it-s} \\ &+ \sum_{m=0}^{M}\hat{\alpha}_{4m}\Delta\ \log\,(1+long)_{it-m} \\ &+ \hat{\alpha}_5\mathrm{ecm}_{it-1} + \varepsilon_t. \end{aligned} \tag{14}$$

The error term has to be included at lag 1 because the *change* in house prices from log hp_{it-1} to log hp_{it} responds to a non-zero error term in $t-1$. The other exogeneous variables can be included in any combination of different lag settings. Table 2 shows the error correction model for the full model including all 15 countries as well as for group 1 and group 2.

Except for group 1, which is displayed only for completeness, the error correction term has the right sign and is highly significant. The value of the error term of around −0.04 indicates that deviations from equilibrium are quite persistent. In particular, half of the equilibrium gap remains after about 17 quarters or four years while 56 quarters or 14 years are required for nearly full adjustment (90% adjustment). On an annual basis, the value of the error term is just −0.16. This value is below previous findings, which ranged from −0.34 to −0.83 (see Hendry, 1984; Drake, 1993; Knudsen, 1994; Eitrheim, 1995; Meen, 1996; Kosonen, 1997; Hort, 1998). But this is above the -0.05 value found by Abra-

Table 2
Error correction model for the adjustment process of international house prices. Group 1 is estimated by fixed effects, whereas the full model and group 2 use random effects with Swamy and Arora (1972) transformation. t-statistics in brackets are computed using Arellano (1987) clustered standard errors. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Variable	Full Model N = 15	Group 1 N = 5	Group 2 N = 9
Constant	0.001^{*} (1.702)	–	0.001^{*} (1.788)
Δ log hp (−1)	0.271^{***} (3.637)	0.600^{***} (5.876)	0.251^{***} (3.066)
Δ log hp (−3)	0.129^{***} (3.552)	-0.243^{***} (−6.108)	0.148^{***} (3.941)
Δ log hp (−4)	0.286^{***} (5.094)	0.599^{***} (17.160)	0.274^{***} (4.553)
Δ log hp (−5)	−0.052 (−1.296)	-0.217^{***} (−3.329)	−0.039 (−0.904)
Δ log ea	0.323^{***} (4.096)	0.087^{***} (49.448)	0.347^{***} (4.297)
Δ log(1 + long)	-0.033^{**} (−2.301)	-0.077^{**} (−2.177)	-0.035^{**} (−2.461)
Δ log(constr)	0.069^{***} (3.475)	-0.038^{***} (−2.710)	0.068^{***} (3.060)
ecm (−1)	-0.040^{***} (−8.277)	0.012^{***} (7.173)	-0.041^{***} (−6.846)
Hausman test	9.051	13.797^{*}	11.858

ham and Hendershott (1996) for the U.S. market. The long adjustment periods are probably attributable to the general downward price stickiness in residential housing. One might expect the error term to be underestimated due to the inclusion of the housing bubble period starting after 2000. The data period includes a large deviation from equilibrium without the subsequent adjustment. However, when reestimating the model from 1975Q1 to 1999Q4 the error term becomes insignificant.[25] Thus, different observation periods are likely to explain a significant part in the differences between the values found in previous research but the inclusion of the bubble period does not lead to systematic underestimation. Furthermore, in a recent study Caballero and Engel (2003) show that the speed of adjustment for a sluggish variable such as house prices may be strongly overestimated when dealing with data at low levels of aggregation. Due to this fact and in face of the larger number of observations of our study we believe our results to be more representative and about fourteen years to be a reasonable time period for nearly full adjustment.[26, 27]

Finally, note also that all other variables have the right sign and are in line with previous findings. For instance, the significant structure of autocorrelation with negative values at lag five has also been found by Englund and Ioannides (1997) who compare the dynamics of house prices in 15 OECD countries. The elasticities for economic activity, interest rates and construction costs are generally smaller than their long-term DOLS results. This supports the view that because of house price stickiness macroeconomic effects add up over time so that an analysis of the long-term and the short-term is necessary.

[25] The data passed both tests of non-stationarity and cointegration. The error term was estimated to be −0.008 (0.127).

[26] The coefficients have been estimated with fixed and random effects. In the presence of lagged endogenous variables these estimators are biased. In this case dynamic panel estimators such as the Arellano and Bond (1991) or the Blundell and Bond (1998) system-GMM estimator yield unbiased results. For macro panels with large time dimension T, however, it has become practice to use fixed effects or random effects since the bias is rather small but the variance of dynamic panel estimators is quite large.

[27] Computing the error term as in (13) leads to significant loss in degrees of freedom since a missing value in only one regressor returns a missing value for the whole error term. For the sake of robustness we therefore report only the adjustment coefficient $\hat{\alpha}_5$ for the whole panel instead of individual country coefficients $\hat{\alpha}_{5i}$.

4. Conclusion

This study examines the impact of the macroeconomy on house prices. Housing market data is often difficult to obtain or covers only short time periods. Using a panel of 15 countries over a period of over 30 years allows for the robust estimation of long-term macroeconomic impacts.

In this context standard theoretical equilibrium models are clearly supported by the empirical results and suggest that macroeconomic variables significantly impact house prices. In particular, a 1% increase in economic activity raises the demand for houses and thus house prices over the long run by 0.6%. An increase in construction costs has an average long-term impact of 0.6% on house prices by reducing housing supply, which leads to an increase in rents and thus in house prices. Finally, an increase in the long-term interest rate makes other fixed-income assets more attractive relative to residential property investment, reducing the demand for this kind of investment which in turn lowers house prices by 0.3% in the long run.

We find weak evidence of an international housing market result in the sense that a selected group of nine countries show a similar long-run response to macroeconomic changes. This may ultimately be useful for predicting the long-term tendencies of the global housing market in the presence of global macroeconomic shocks. In contrast to current literature, our findings suggest that the speed of adjustment to equilibrium may be actually much slower than has been previously suggested. While 14 years may at first appear rather long for nearly full adjustment, we believe it is a reasonable time frame given the stickiness in residential house prices. The size and the panel structure of our dataset allow us to gain more robust results. Furthermore, there is evidence that the speed of adjustment may be strongly overestimated when dealing with low levels of aggregation so that our findings are probably closer to the true values. This is also supported by the fact that the coefficients on all other variables including those of the rich autoregressive structure are in line with previous research.

Further research should take possible endogeneity between the rate of adjustment in house prices and construction costs into account. As house prices adjust upwards (downwards) construction costs eventually increase (decrease) as well adding to the effect of high (low) demand.

Modeling this relationship explicitly may provide additional insights into the underlying structure of house price adjustments in international housing markets.

Acknowledgments

We would like to thank the editor Henry Pollakowski, two anonymous referees, Bernd Fitzenberger, Reint Gropp, Peter Pedroni, Joachim Zietz as well as the participants of the 2008 Conference "What Drives Asset and Housing Markets?" at the Center for European Economic Research (ZEW) for useful comments and suggestions. We thank Tiana Moren-Jefferies from the Bank for International Settlements for providing us with the housing market data. We alone are responsible for any errors.

References

Abraham, J., Hendershott, P., 1996. Bubbles in metropolitan housing markets. Journal of Housing Research 7 (2), 191–207.

Annett, A., 2005. Euro area policies: selected issues. IMF Country Report No. 05/266.

Arellano, M., 1987. Computing robust standard errors for within group estimators. Oxford Bulletin of Economics and Statistics 49 (4), 431–434.

Arellano, M., Bond, S., 1991. Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations. Review of Economic Studies 58 (2), 277–297.

Berg, L., 1996. Age distribution, saving and consumption in Sweden, Uppsala University Working Paper 1996:22. Available from SSRN: <http://ssrn.com/abstract=1405> or doi: 10.2139/ssrn.1405.

Berg, L., Bergström, R., 1995. Housing and financial wealth, financial deregulation, and consumption—the Swedish case. Scandinavian Journal of Economics 97, 421–439.

Blundell, R., Bond, S., 1998. Initial conditions and moment restrictions in dynamic panel data models. Journal of Econometrics 87 (1), 115–143.

Bowen, A., 1994. Housing and the macroeconomy in the United Kingdom. Housing Policy Debate 5 (3), 241–251.

Breitung, J., Das, S., 2005. Panel unit root tests under cross sectional dependence. Statistica Neerlandica 59 (4), 414–433.

Caballero, R., Engel, E., 2003. Adjustment is Much Slower than You Think, Cowles Foundation Discussion Paper No. 1430. Available from SSRN: <http://ssrn.com/abstract=430040>.

Campbell, J.Y., Cocco, J.F., 2007. How do house prices affect consumption? Evidence from micro data. Journal of Monetary Economics 54, 591–621.

Case, K.E., 2000. Real estate and the macroeconomy. Brookings Papers on Economic Activity 20 (2), 119–162.

Case, K.E., Quigley, J.M., Shiller, R.J., 2005. Comparing wealth effects: the stock market versus the housing market. Advances in Macroeconomics 5 (1), 1–32.

Catte, P., Girouard, N., Price, R., André, C., 2004. Housing markets, wealth and the business cycle. OECD Working Paper No. 394.

Cocco, J.F., 2005. Portfolio choice in the presence of housing. Review of Financial Studies 18 (2), 535–567.

DiPasquale, D., Wheaton, W.C., 1996. Urban Economics and Real Estate Markets. Prentice-Hall, Englewood Cliffs, New Jersey.

Drake, L., 1993. Modelling UK house prices using cointegration: an application of the Johansen technique. Applied Econometrics 25 (9), 1225–1238.

Eitrheim, Ø., 1995. Norwegian Housing Prices 1983 to 1992: A Linear Dynamic Model. Mimeo, Central Bank of Norway.

Engelhardt, G.V., Poterba, J.M., 1991. House prices and demographic change: Canadian evidence. Regional Science and Urban Economics 21 (4), 539–546.

Engle, R.F., Granger, C.W.J., 1987. Co-integration and error-correction: representation, estimation and testing. Econometrica 55 (5), 251–276.

Englund, P., Ioannides, Y., 1997. House price dynamics: an international empirical perspective. Journal of Housing Economics 6 (2), 119–136.

Girouard, N., Kennedy, M., van den Noord, P., André, C., 2006. Recent house price developments: the role of fundamentals. OECD Working Paper No. 475.

Glaeser, L.E., 2000. Comments and Discussion on: K.E. Case (2000), Real estate and the macroeconomy. Brookings Papers on Economic Activity 20 (2), 146–150.

Hanck, C., 2008. An intersection test for panel unit roots. Working Paper. University of Dortmund.

Heiborn, M., 1994. Demographic factors and the demand for housing. Working Paper. Uppsala University.

Hendry, D.F., 1984. Econometric modelling of house prices in the United Kingdom. In: Hendry, D.F., Wallis, K.F. (Eds.), Econometrics and Quantitative Economics. Blackwell Publishers, pp. 211–252.

Hort, K., 1998. The determinants of urban house price fluctuations in Sweden 1968–1994. Journal of Housing Economics 7 (2), 93–120.

Hunt, B., Badia, M., 2005. United Kingdom: selected issues. IMF Country Report No. 05/81.

Im, K., Pesaran, H., Shin, Y., 1997. Testing for unit roots in heterogeneous panels. Discussion Paper. University of Cambridge. June.

Inder, B., 1993. Estimating long-run relationships in economics: a comparison of different approaches. Journal of Econometrics 57 (1–3), 53–68.

Johansen, S., Juselius, K., 1990. Maximum likelihood estimation and inference: on cointegration with applications to the demand for money. Oxford Bulletin of Economics and Statistics 52 (2), 169–210.

Kao, C., Chiang. M.H., 2000. On the estimation and inference of a cointegrated regression in panel data. In: Baltagi, B.H. (Ed.), Advances in Econometrics: Nonstationary Panels, Panel Cointegration and Dynamic Panels, vol. 15. Elsevier Science Inc., pp. 179–222.

Knudsen, D., 1994. Residential investments and house prices in Denmark. Economic Modelling 11 (2), 201–214.

Koskela, E., Loikkanen, H., Virén, M., 1992. House prices, household saving, and financial market liberalization in Finland. European Economic Review 36, 549–558.

Kosonen, K., 1997. House price dynamics in Finland. A time series analysis of the period 1979–95. Labour Institute for Economic Research. Discussion Papers 137.

Leung, C., 2004. Macroeconomics and housing: a review of the literature. Journal of Housing Economics 13 (4), 249–267.

Levin, A., Lin, C., 1993. Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties, University of California, San Diego Discussion Paper, December.

Maddala, G.S., Wu, S., 1999. A comparative study of unit root tests with panel data and a new simple test. Oxford Bulletin of Economics and Statistics 61, 631–652.

Mankiw, G.N., Weil, D.N., 1989. The baby boom, the baby bust, and the housing market. Regional Science and Urban Economics 19 (2), 235–258.

Mark, N.C., Sul, D., 2003. Cointegration vector estimation by panel DOLS and long-run money demand. Oxford Bulletin of Economics and Statistics 65 (5), 655–680.

McCarthy, J., Peach, R., 2004. Are home prices the next bubble? Federal Reserve Bank of New York Economic Policy Review, Federal Reserve Bank of New York, December.

Meen, G., 1996. Ten propositions in UK housing macroeconomics: an overview of the 1980s and early 1990s. Urban Studies 33 (3), 425–444.

Meen, G., 2002. The time-series behavior of house prices: a transatlantic divide. Journal of Housing Economics 11 (1), 1–23.

Moon, H.R., Perron, B., 2004. Testing for a unit root in panels with dynamic factors. Journal of Econometrics 122 (1), 81–126.

Parker, J.A., 2000. Comments and discussion on: K.E. Case (2000), Real estate and the macroeconomy. Brookings Papers on Economic Activity 20 (2), 150–158.

Pedroni, P., 1999. Critical values for cointegration tests in heterogeneous panels with multiple regressors. Oxford Bulletin of Economics and Statistics 61 (4), 653–670.

Pedroni, P., 2000. Fully modified OLS for heterogeneous cointegrated panels. Advances in Econometrics 15, 93–130.

Pedroni, P., 2004. Panel cointegration; asymptotic and finite sample properties of pooled time series tests, with an application to the PPP hypothesis. Econometric Theory 20 (3), 597–625.

Phillips, P.C.B., Moon, H., 1999. Linear regression limit theory for nonstationary panel data. Econometrica 67 (5), 1057–1111.

Phillips, P.C.B., Sul, D., 2003. Dynamic panel estimation and homogeneity testing under cross section dependence. Econometrics Journal 6 (1), 217–259.

Poterba, J., 1991. House price dynamics. Brookings Papers on Economic Activity 20 (2), 143–203.

Quah, D., 1994. Exploiting cross-section variation for unit root inference in dynamic data. Economics Letters 44 (1), 9–19.

Saikkonen, P., 1991. Asymptotically efficient estimation of cointegration regressions. Econometric Theory 7 (1), 1–21.

Stock, J.H., Watson, M.W., 1993. A simple estimator of cointegrating vectors in higher order integrated systems. Econometrica 61 (4), 783–820.

Sutton, G., 2002. Explaining changes in house prices. BIS Quarterly Review, Bank for International Settlements, September.

Swamy, P., Arora, S., 1972. The exact finite sample properties of the estimators of coefficients in the error components regression models. Econometrica 40 (2), 261–275.

Terrones, M., Otrok, C., 2004. The global house price boom. IMF World Economic Outlook, September.

Tsatsaronis, K., Zhu, H., 2004. What drives housing price dynamics: cross country evidence. BIS Quarterly Review, Bank for International Settlements, March.

Wagner, R., 2005. En Model for de Danske Ejerboligpriser, Okonomi – og Erhvervsministeriets arbejdspapir, No. 1.

Zuh, H., 2003. The importance of property markets for monetary policy and financial stability. BIS Working Paper 21.