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What Drives Housing Prices Down? Evidence from an International Panel

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What drives housing prices down? Evidence from an international panel §

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December 2007

Abstract

In this study, we suggest an explanation for the alarmingly low growth rates of real housing prices in Canada and Germany in comparison to other OECD countries over 1975-2005. We show that the long-run development of housing markets is determined by real disposable per capita income, real long-term interest rate, population growth, and urbanization. The differential development of real housing prices in Canada and Germany is attributed to the specific values of the fundamentals in these two countries. Canada and Germany are characterized by relatively low average growth rates of real disposable income and relatively high interest rates resulting in suppressed housing prices over long period of time. Institutional structure accentuates these tendencies. Given the importance of housing wealth for the private consumption, our paper aims at drawing attention of the policymakers to the necessity of preventing not only the overheating but also overcooling of the housing market that entails lower economic growth rate.

Keywords: house prices; dynamic panel data; cointegration.

JEL classification: E30; C23; C51.

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1 Introduction

The last few years saw a very fast increase in the housing prices in many countries. In some countries such as Ireland, Spain, UK, and USA, the growth rates of housing prices were so alarmingly high that it raised fears about emerging speculative bubbles. The worries are reinforced by the recent US sub-prime mortgage crisis, which has lead to plunging property prices and a slowdown in the US economy. It is feared that the "US scenario" can repeat in other countries with booming housing prices. However, these discussions neglect another group of countries, where real house prices have been stagnating and even decreasing over the last few decades.

This diverging house price[1] development among the OECD countries is illustrated in Figure 1 that shows the real house price (nominal house prices net of consumer price inflation) dynamics in 14 OECD countries in 1975-2004. These countries are classified in four groups: a) countries with falling house prices (Germany and Canada — upper left panel); b) countries with stagnating house prices (Japan and Switzerland — upper right panel); c) countries with a medium house price growth (Belgium, Finland, France, Italy, Sweden, and USA — lower left panel), and finally d) countries with an extremely high house price growth (Ireland, the Netherlands, Spain, and the UK — lower right panel). As can be seen, over the last 30 years, the real house prices in Germany and Canada experienced almost uninterrupted decline. Japanese and Swiss house prices grew up to the late 1980s — early 1990s and then started to decrease. In Switzerland a decade later, this decline turned into an upswing, whereas in Japan the real house prices continued to fall. In contrast, the real house prices in other countries in our sample followed an upward trend, sometimes interrupted for short periods of time.

The objective of this paper is thus to investigate the factors, which prevent real housing prices from rising or even drive them down for a protracted period of time. House price development is important from the policymaker perspective, since it can affect consumers' wealth and hence private consumption, see for example Carroll et al. (2006) and Slačálek (2006). As the recent German experience shows, an under-consumption caused by the negative house wealth effect may set the economy on a lower equilibrium path.

Until now little attention has been paid to the adverse effects of stagnating or falling house price. Virtually all discussions have been concentrated upon the dangerous consequences of bursting

[1] See Table 1 for source of the data.

speculative bubbles at the housing market. Our paper aims at drawing attention of the policymakers to the necessity of preventing not only the overheating but also overcooling of the housing market.

Our paper contributes to a small body of literature, which analyzes determinants of housing prices based on a panel of countries or regions. To the best of our knowledge, there exist only five studies of housing price determinants, which employ panel data. These include: Almeida et al. (2006), Annett (2005), Égert and Mihaljek (2007), Terrones and Otrok (2004), which use international data, and Holly et al. (2007), which concentrate on the USA states. A detailed overview of these studies is given in Table 3[2].

Whereas Terrones and Otrok (2004), Almeida et al. (2006), and Annett (2005) estimate pooled panel-data models, where all slope parameters are restricted to be the same across countries, Égert and Mihaljek (2007) and Holly et al. (2007) employ mean group estimators based on the averages of the individual estimates made for each country separately. The mean group estimators applied by Égert and Mihaljek (2007) and Holly et al. (2007), however, do not take into account the fact that certain parameters may be identical across individuals (countries).

In what follows we apply the pooled mean group estimator of Pesaran et al. (1999). To the best of our knowledge, this is the first time the pooled mean group estimator is used to analyze the determinants of housing prices. This estimator imposes equal long-run parameters and country-specific short-run parameters and error variances. It can thus be considered as an intermediate case between the pooled and mean group estimators, since it involves both pooling and averaging. The assumption of common long-run parameters can be considered as a rather restrictive. However, it seems warranted, given that our analysis includes a relatively homogeneous group of industrialized countries and that allowing for country-specific short-run parameters captures all the relevant heterogeneity across countries.

The paper is structured as follows. In section 2, we describe the standard housing price determinants suggested in the literature as well as the data, which we use in this study. In section 3 our methodology is presented and estimation and specification tests' results are discussed. Section 4 provides an interpretation of our econometric results. Section 5 concludes.

[2] The explanation of abbreviated country codes used in this table and elsewhere in the text can be found in Table 2.

2 Data

2.1 Determinants of house prices

The factors affecting housing prices may be classified into three major categories: economic factors (income, interest rate, credit, stock market), demographic factors (population growth, urbanization, household size), and institutional factors (financial system, taxation system).

An excellent survey of the current literature on determinants of house prices is contained in a recent study by Girouard et al. (2006). According to this survey, the following three economic variables are the most important for the understanding of the house price development: on the supply side, existing housing stock, which is negatively related with house prices; on the demand side, real disposable income per capita, which positively affects house prices, and real long-term interest rate, which negatively affects house prices.

According to Girouard et al. (2006), demographic determinants of house prices typically include: net migration, total population, and population growth. However, results concerning the demographic variables are rather mixed. For instance, some studies find the population growth to be an important determinant of house prices, whereas other find it to be insignificant.

Finally, house prices can be affected by such institutional factors as taxation policies, financial systems, etc. However, institutions are very difficult to measure and, if we are not considering the economies in transition (see Égert and Mihaljek (2007)), change very slowly. Thus, over relatively short samples (30 years and less), they remain basically unchanged.

2.2 Our data

Our data set includes an unbalanced panel of 14 OECD countries[3] (Belgium, Canada, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Spain, Sweden, Switzerland, UK, and the USA) and covers the period 1974-2005. The sources and transformations of the data used in this study are described in Table 1.

Data on the housing prices were obtained from the NIGEM-model, which is based on a data set delivered by the Bank for International Settlements (BIS). BIS collects price indices not only from national central banks and statistical offices, but also from commercial sources. The house price indices, partly, differ in terms of assessment bases and types of dwellings: National indices are

[3]The choice of countries is determined by the availability of the data.

mostly calculated based on regional data, which include purchase prices from newly built houses as well as from resales. Real house prices are obtained through deflating the nominal house prices with the consumer price index.[4]

The explanatory variables used in this study include both economic factors and demographic ones. The economic variables are the real disposable income per capita and real long-term interest rate. The last variable deserves a special attention. In our study the nominal long-term interest rate is deflated using the housing price index and not consumer price index (CPI) as in some other studies, e.g., in Annett (2005). Our definition of real interest rate is based on the reasoning of a household, which makes decision about buying an housing asset. It compares the income it can earn on a bank deposit with a "capital gain" stemming from changes in housing prices. The long-term interest rate is chosen because buying a house or an apartment is a long-term investment. The demographic variables are population growth and urbanization degree.

3 Estimation and testing

3.1 Model specification

We employ the pooled mean group (PMG) estimator suggested in Pesaran et al. (1999) to our data. The model specification in the error-correction form is as follows:

$$\begin{aligned} \Delta RHP_{it} = {} & \phi_i(RHP_{i,t-1} - \theta_{i0} - \theta_1 RDI_{it} - \theta_2 RLIR_{i,t-1} - \theta_3 POP_{it} - \theta_4 URB_{it}) + \sum_{h=1}^{p-1} \gamma_{ih} \Delta RHP_{i,t-h} \\ & + \sum_{j=0}^{q-1} \psi_{ij}^1 \Delta RDI_{it-j} + \sum_{j=1}^{q} \psi_{ij}^2 \Delta RLIR_{i,t-j} + \sum_{j=0}^{q-1} \psi_{ij}^3 \Delta POP_{it-j} + \sum_{j=0}^{q-1} \psi_{ij}^4 \Delta URB_{it-j} + \varepsilon_{it} \end{aligned} \quad (1)$$

where the common values are imposed on the long-run coefficients and where RHP_{it} is the real house price (nominal house price net of consumer price inflation); RDI_{it} is the real disposable income per capita; $RLIR_{it}$ is the real long-term interest rate (nominal interest rate minus growth rate of the house prices); POP_{it} is the population growth, and URB_{it} is the urbanization degree. Notice that in order to avoid the simultaneity problem, the real long-term interest rate in equation (2) is taken with lag. Observe that the error-correction model can be trivially derived from the

[4]We refer to Appendix for a detailed description of all data used.

following Auto-Regressive Distributed Lag model, $ARDL(p,q,q,q,q)$:

$$RHP_{it}=\sum_{h=1}^{p}\lambda_{ij}RHP_{i,t-j}+\sum_{j=1}^{q}\delta_{ij}^{1}RDI_{i,t-j}+\sum_{j=1}^{q+1}\delta_{ij}^{2}RLIR_{i,t-j}+ \\ +\sum_{j=0}^{q}\delta_{ij}^{3}POP_{i,t-j}\ +\sum_{j=0}^{q}\delta_{ij}^{4}URB_{i,t-j}+\mu_i+\varepsilon_{it} \quad (2)$$

Given the fact that we have annual time series with a maximum length of up to 32 observations, we impose $p = q = 2$. Our subsequent analysis of residuals shows such lag augmentation structure is fully sufficient to remove autocorrelation from model residuals. We also allow for differential lag augmentations as the optimal lag length in every cross-section is chosen by minimizing the Akaike information criterion.

3.2 Estimation results

The estimation results of the model are presented in Table 7. Here we focus only on the long-run parameters values, values of the adjustment coefficients of the error correction term, half-life of shock, which measures the time necessary for a deviation from long-run equilibrium to be halved (half-life=$\frac{ln(0.5)}{ln(1+\phi_i)}$), and the measure of goodness of fit of our empirical regressions.

First, notice that the point estimate of long-run elasticity of real house prices with respect to income is very close to unity and according to the reported standard error is also insignificantly different from that value. This finding is consistent with the theoretical considerations discussed in Holly et al. (2007). The effect of the real long-term interest rate on real house prices is found to be significantly negative as it measures foregone return on the alternative assets compared to return on housing. Furthermore, the other two explanatory variables such as population growth and degree of urbanization, as expected, positively influence real house prices.

Second, the heterogeneous estimates of the adjustment coefficient ϕ_i are negative for all countries and are significantly different from zero at least at the 5% level in eleven out of 14 cases. The reported negative sign of the adjustment coefficients suggests that correction of the past disequilibria indeed takes place, albeit its adjustment speed varies from country to country. As seen, for most countries the past disequilibrium is half-corrected in less than ten years. However, there are certain exceptions like Japan and Switzerland were reported adjustment coefficients are very close to zero and appear to be statistically insignificant from zero. This has to be traced to the bust of house price bubble observed in both countries around 1990.

Third, our model is able to explain up to 80% of variation in the real house prices, depending on a country. The adjusted R^2 values vary from 0.255 for Italy to 0.839 for the UK.

3.3 Robustness check

All in all, the estimation results presented in the section above point out that our long-run parameter estimates are sensible and, moreover, adjustment towards the long-run equilibrium takes place in every country. In this subsection, we check the robustness of our results by estimating parameters of the dynamic panel data model using the same specification as above but omitting one country at a time. The results of this exercise are displayed in Figures 2 and 3, where the boxplot of the adjustment coefficient values is presented as well as the sequence of the long-run parameter estimates, respectively.

In Figure 2 we observe that the distribution of adjustment coefficient values undergoes quite little variation regardless what country has been omitted from our sample, with a slight exception for France where the interquartile range seems somewhat smaller and more skewed towards zero than in the rest of cases. Nevertheless, the robustness of our results is supported by the fact that all estimated adjustment coefficients calculated for all combinations of countries have negative sign with the median fluctuating around -0.1.

The results of a similar exercise concerning the long-run parameter estimates are reported in Figure 3. Also there we observe rather remarkable stability of estimates with one exception. In a panel that excludes France the estimate of real interest rate seems to be almost twice as large as those reported when any other country was omitted. At the same time, omission of France resulted in somewhat lower estimate of real income effect. However, it is still statistically insignificantly different from unity, based on the reported 95% confidence interval. As regarding the other two variables — population growth and urbanization — omitting France from a panel seems to be little noticeable. Thus, the overall impression is that both adjustment coefficients and long-run coefficients display rather high degree of robustness with respect to omitting a single country at a time and the outcome with France could be safely seen as a separate incident. This implies that our main estimation results reported in Table 7 are not due to some outlying observations associated with a particular country.

3.4 Non-stationarity and cointegration

In this subsection, we address important issues about which a researcher should be concerned when dealing with the persistent time series. First, we discuss testing for order of integration of the variables in question in panel data by allowing for cross-sectional interdependence. Second, we will test for cointegration in our model.

In the panel data literature, there was suggested a number of panel unit-root tests like Maddala and Wu (1999); Hadri (2000); Levin et al. (2002); Im et al. (2003), *inter alia*. However, as pointed out in Strauss and Yigit (2003) and Jönsson (2005) these tests are based on a rather restrictive assumption of cross-sectional independence, which in reality is very likely to be violated, given economic, political, cultural, and other linkages between different economies. As a consequence, these generation of panel unit-root tests has poor size properties and low power in the presence of cross-sectional dependence.

In our paper, in order to circumvent the problem of cross-sectional dependence when testing for unit roots we proceed in two steps. First, we apply a cross-sectional dependence (CD) test suggested in Pesaran (2004) where the corresponding null hypothesis is that cross-sectional dependence is absent in our data. This test is very simple to implement since it is based on an average of the pairwise correlations of the OLS residuals obtained from the individual regressions in the panel. The CD test statistic has a standard normal limiting distribution. At the second step, depending on the outcome of the CD test, we employ either tests that are based on the assumption of no cross-sectional dependence or the tests suggested in Moon and Perron (2004) and Pesaran (2007) that are robust to the presence of cross-sectional dependence.

Table 4 presents the results of the CD test of Pesaran (2004). The entries in the table are the corresponding test statistics computed using the residuals of the Augmented Dickey-Fuller (ADF) unit-root test regressions with maximum lag length augmentation order $k = 0, 1, 2, 3, 4$. The ADF test was applied to every cross-sectional unit $i = 1, 2, \ldots, 14$ of every variable. The CD test decisively rejects the null hypothesis of absence of cross-sectional dependence in the residuals of the individual ADF regressions, implying that panel unit-root tests that do not account for cross-sectional dependence are inappropriate in these circumstances.

In sequel, we apply the panel unit-root tests of Pesaran (2007) and Moon and Perron (2004) that are designed to tackle cross-sectional dependence in the tested time series. The former test is based on the individual ADF regressions augmented with cross-section averages (henceforth, CADF) in order to filter out cross-sectional dependence in the regression residuals. Since this test is a direct generalization of the panel unit-root test of Im et al. (2003), the corresponding test statistics is denoted as a cross section IPS (CIPS) and is based on a simple average of individual CADF t-ratios $\widetilde{t}_i$

$$\text{CIPS} = \frac{1}{N}\sum_{i=1}^{N}\widetilde{t}_i \quad (3)$$

The latter test explicitly models cross-sectional dependence in the data by allowing for up to m common factors in the panel, which are estimated using the principal component analysis. The test statistic t_b^* has a limiting standard normal distribution as both $N \to \infty$ and $T \to \infty$ as long as $N/T \to 0$.

Tables 5 and 6 present the results of Pesaran (2007) and Moon and Perron (2004) panel unit-root tests allowing for cross-sectional dependence, respectively. We apply these tests both to the levels and to the first differences of the time series. The results of these two tests largerly imply that the real house price RHP_{it}, the real disposable income RDI_{it}, and population growth POP_{it} are I(1) variables, whereas the real interest rate $RLIR_{it}$ and URB_{it} are I(0). Although the test outcomes are somewhat sensitive to specified lag augmentation p of the individual CADF tests in case of Pesaran (2007) test and to various number of factors m in case of Moon and Perron (2004). Testing for order of integration of the first-difference transformation of our variables in question uniformly leads to a conclusion that they are stationary according to Moon and Perron (2004) test, whereas conclusions based on the test of Pesaran (2007) vary with the length of lag augmentation in the auxiliary unit root test regressions p.

Provided that some of our variables appear to be non-stationary, it is necessary to address the issue of existence of cointegration in order to rule out the possibility of spurious regression. Unfortunately, to the best of our knowledge, there is no direct test for cointegration that can be applied to the PMG model (1). Therefore, we chose to test for existence of a cointegrating relationship in our data set using an indirect method. Notice that existence of cointegration implies that the adjustment coefficients of the error-correction term should be significantly different from zero and have a correct sign, so that the past deviations from the implied equilibrium are partially eliminated in the current period. As seen from Table 7, in all cases the adjustment coefficient of the error correction term has an expected negative sign and it is significantly different from zero at least at the 5% level in eleven and at the 1% level in nine out of 14 cases. This strongly suggests existence of a cointegration relationship (augmented with the I(0) variables) that forms our error-correction mechanism[5].

4 Interpretation of results

In this section, we address the main question of this paper: What drives the real housing prices down? In particular, we concentrate on two countries (Canada and Germany) with falling real housing prices, which stay in a sharp contrast with most countries in our sample.

As our estimation results in section 3 show, the following economic and demographic factors determine the long-run development of real housing prices: real disposable income, real long-term interest rate, population growth, and urbanization. While real disposable income, interest rate, population growth, and urbanization exert a positive impact on the housing price dynamics, real long-term interest rate dampens the growth of real housing prices. In addition, the long-run coefficient at real disposable income is very close to unity, which implies that in the long run real housing prices move in line with real income conditioned upon other factors. Our results thus conform with implications of the theory — see Holly et al. (2007). In this respect our study differs from those of Annett (2005) and Égert and Mihaljek (2007), who report considerably lower than unity estimates of the long-run income elasticity. At the same time, our results are close to those obtained by Almeida et al. (2006) and Holly et al. (2007).

Furthermore, our model implies that the long-run influence of fundamentals on real housing prices is very similar across the different countries. Therefore, a question arises why in some countries real housing prices continue to fall for a protracted period of time whereas in other countries housing market is booming. A deeper analysis of fundamentals is necessary to answer this question.

Figure 4 reports average values of explanatory variables in our model over the period 1975-2004 for individual countries. We sorted these average values in ascending order. A closer look at the figure reveals that Canada and Germany are similar in many respects. Thus, these countries show low growth rates of real disposable income per capita and the highest real long-run interest rates.

[5] Égert and Mihaljek (2007) also do not conduct a formal test for cointegration but instead argue that negative and significant adjustment coefficient provides a sufficient evidence for existence of a cointegrating relationship in their empirical model.

Both these countries are characterized by relatively low growth rates of urbanization. The only respect in which Canada and Germany differ is the population growth: Canada is a country with fastest population growth, whereas Germany ranks near the bottom of the country list. One can safely conclude that it is the fundamentals that curb Canadian and German housing markets and make them so different from housing markets of other countries.

The limitation of current study is that it exclusively addressed influence of economic and demographic factors on housing markets. At the same time, influence of institutional factors, which cannot be overstated, was not taken directly into account due to the fact that institutional arrangements are difficult to measure and they tend to change very slowly in time. Following the standard practice of panel data literature, their influence was indirectly captured by fixed effects as well as via country-specific short-run coefficients, which are allowed to in our modelling approach. ECB (2003) provides a comprehensive analysis of institutional characteristics of major European housing markets. This study shows that in general Germany is characterized by relatively stricter taxation policy, discouraging speculations with real estate, less controlled rental markets and less subsidized home-ownership such that the share of rented dwellings stayed well above 50%, while it dropped considerably in most of the other European countries over the period 1980–2000 (see ECB (2003), p. 26). In addition, Germany remains one of the countries with the highest degree in housing market regulation, according to SVR (2006).

Similarly, one can conclude that the particular conditions in the Canadian financial system have also contributed to stagnating housing prices over the most period examined in this study, except for the last few years when Canadian housing market began its modest recovery. Even if certain restrictions on banks with respect to their mortgage financing have been abolished in 1967 (Girouard and Blöndal (2001)), the lending behavior of banks remained conservative: "The mortgage credit culture in Canada is rather conservative, with a large majority of mortgages at fixed interest rates and a preference for mortgage terms of five years. Interest bearing term instruments sold to savers remain the primary source of funding for mortgage loans, which subsequently remain largely on the balance sheet of lenders" (see Traclet (2005, p.1)).

5 Conclusion

In this study, we suggest the following explanation for the extraordinary weak development in Canadian and German real housing prices in comparison to other industrialized countries. As our econometric analysis shows, the general long-run development of housing markets in these countries is equally determined by such factors as real disposable per capita income, real long-term interest rate as well as by population growth and urbanization.

The differential development of real housing prices in Canada and Germany can be attributed to the specific values of the fundamentals in these two countries. Canada and Germany are characterized by relatively low average growth rates of real disposable income and relatively high interest rates resulting in suppressed housing prices over long period of time. Institutional structure accentuates these tendencies.

Given the importance of housing wealth for the private consumption, our paper aims at drawing attention of the policymakers to the necessity of preventing not only the overheating but also overcooling of the housing market that entails lower economic growth rate.

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Appendix

Table 1: Data sources

Variable	**Notes**	**Source**	**Code**
Nominal house price index	1995 = 100	NIGEM, based on BIS Ireland and Switzerland taken directely from BIS	HP
Consumer price index	all components, 2000 = 100	OECD, Main Economic Indicator	CPI
Real house price index		$\ln(HP/CPI)$	RHP
Real disposable income	national currencies	AMECO	RDI_TOT
Real disposable income per capita		$\ln(RDI\_TOT/POPUL)$	RDI
Nominal long-term interest rate		OECD Economic Outlook, No. 80 For Ireland, money market rate, International Monetary Fund (IMF)	LIR
Real long-term interest rate		$LIR - \Delta\ln(HP)$	RLIR
Population	millions of persons	AMECO, for Germany, Federal Statistical Office	POPUL
Population growth		$\ln(POPUL_t/POPUL_{t-1})$	POP
Urbanization degree	log of share of urban population	World Bank, World Market Indicator	URB

Table 2: Country codes used in this paper

Code	Country	Code	Country
AU	Australia	KR	South Korea
BE	Belgium	MY	Malaysia
CA	Canada	NL	Netherlands
CL	Chile	NZ	New Zealand
DK	Denmark	NO	Norway
FI	Finland	SG	Singapore
FR	France	ES	Spain
DE	Germany	SE	Sweden
HK	Hongkong	CH	Switzerland
IE	Ireland	TW	Taiwan
IL	Israel	TH	Thailand
IT	Italy	UK	UK
JA	Japan	US	USA

Table 3: Methodological approaches

Study	Annett (2005)	Almeida et al. (2006)[2]	Terrones and Otrok (2004)	Holly et al. (2007)[4]	Égert and Mihaljek (2007)
Countries	DE, FR, IT, IE, ES, NL, BE, FI	AU, BE, CA, CL, DK, FI, FR, DE, HK, IE, IL, IT, JA, KR, MY, NL, NZ, NO, SG, ES, SE, CH, TW, TH, UK, US	AU, BE, CA, DK, FI, FR, DE, IE, IT, JA, NL, NZ, NO, ES, SE, CH, US, UK	US States (without Hawaii and Alaska)	8 Central and Easterm Europe and 19 OECD countries
Time period	1970 - 2003	1970 - 1999	1970 - 2003	1975 - 2003	1975-2005
Method	Panel ECM	Pooled panel estimation OLS and Arellano-Bond estimation corrected with White-Hubber estimator	DPD Lamont/Stein (1999) GMM Arellano/Bond (1991)	Panel ECM	Panel DOLS
Explanatory variables					
Income	$+^{***}$	$+^{***}$	$+^{***}$	$+^{***}$	$+^{**}$
Interest rate	$-^{***}$	$-^{**}$	$-^{***}$		$-^{**}$
House prices, lag	$-^{***}$	$+^{***}$	$+^{***}$	0	
Credit	$+^{***}$		$+^{***}$		$+^{**}$
Money	$+^{*}$				
Population growth	0		$+^{***}$	$+^{**}$	$+^{**}$
Labor force					$+^{**}$

Table 3: Methodological approaches (continued)

Study	Annett (2005)	Almeida et al. (2006)[2]	Terrones and Otrok (2004)	Holly et al. (2007)[4]	Égert and Mihaljek (2007)
Stock price index			$+^{***}$		$-^{**}$
Bank crisis			$-^{***}$		
Net cost of borrowing[1]				$-^{**}$	
Lagged housing affordability ratio[3]		$-^{***}$	$-^{***}$		
Term spread					
Inflation		0			
Real wage					$+^{**}$
Error correction	$-^{***}$			$-^{***}$	$-^{**}$
R^2	0.6	0.4			0.65-0.90
Data transformations	variables in log differences, real interest rate in differences	variables in log differences	variables in growth rates, interest in percent	variables in growth rates, with the exception of net cost of borrowing	variables in log differences, real interest rate in differences, population and unemployment as shares
	long-term interest rate, all interest rates deflated with CPI	long-term interest rate	short-term interest rate		short-term interest rate

[1] Holly et al. (2007) define the net cost of borrowing as c = real interest rate - change in real house prices.

[2] Almeida et al. (2006) estimate several submodels with different specifications. Results in this table are taken from their benchmark model, estimated by OLS and without the affordability ratio.

[3] Affordability ratio = ratio of real house prices to (per capita) real income. Beside their benchmark model, Almeida et al. (2006) find a significance negative impact of this ratio, too.

[4] Holly et al. (2007) find cointegration between prices and income and estimate then an ECM.

Table 4: ADF(p) residuals cross section dependence test, 1975-2004

Variable	Lag order				
	0	1	2	3	4
Real house price	13.95	9.37	8.36	7.52	7.1
Real disposable income per capita	22.78	19.2	17.89	17.33	15.6
Real long-run interest rate	11.7	8.9	7.61	6.56	6.63
Population growth	4.52	1.66	1.71	1.74	1.37
Urbanization	4.92	4.58	4.21	3.98	4.04

Notes: Table entries are the CD test statistic of Pesaran (2004) applied to the residuals of the individual ADF(p) test regressions computed for each panel variable. Under the null of no error cross section dependence it has a standard normal limiting distribution.

Table 5: CIPS panel unit-root test, 1975-2004

Variable	CADF(0)	CADF(1)	CADF(2)	CADF(3)	CADF(4)
Levels: with intercept and trend					
RHP_{it}	-1.85	-2.08	-1.38	-1.26	-0.88
RDI_{it}	-1.63	-1.77	-1.53	-1.56	-1.68
$RLIR_{it}$	-3.40***	-3.49***	-2.91**	-2.97***	-2.26
POP_{it}	-1.95	-2.49	-1.68	-1.40	-1.13
URB_{it}	-4.00***	-3.99***	-4.86***	-6.25***	-9.67***
First differences: with intercept					
ΔRHP_{it}	-2.56***	-2.52***	-1.95	-1.95	-1.60
ΔRDI_{it}	-3.48***	-2.77***	-2.22*	-1.79	-1.51
$\Delta RLIR_{it}$	-4.56***	-3.97***	-2.91***	-2.87***	-2.12
ΔPOP_{it}	-3.81***	-3.67***	-2.87***	-2.35**	-1.80
ΔURB_{it}	-2.33**	-2.20*	-2.15*	-1.97	-3.34***

Notes: Table entries are the CIPS(p) statistics of Pesaran (2007). The relevant 10%, 5%, and 1% critical values are -2.66, -2.76, and -2.96 with an intercept and a linear trend case and -2.14, -2.25, and -2.45 with an intercept case. '*', '**', and '***' denote significance at the 10%, 5%, and 1% levels, respectively.

Table 6: Moon and Perron's t_b^* panel unit-root test, 1975-2004

Variable \ No. of factors	1	2	3	4
Levels: with intercept and trend				
RHP_{it}	0.06	-0.11	0.22	-0.35
RDI_{it}	-0.68	-1.00	-4.22***	-3.31***
$RLIR_{it}$	-4.21***	-4.64***	-8.08***	-7.00***
POP_{it}	-0.51	-0.18	-0.04	-1.59*
URB_{it}	-4.07***	-0.45	1.30	0.88
First differences: with intercept				
ΔRHP_{it}	-8.69***	-7.96***	-8.82***	-8.45***
ΔRDI_{it}	-8.90***	-10.37***	-9.97***	-10.85***
$\Delta RLIR_{it}$	-14.96***	-18.07***	-17.09***	-17.99***
ΔPOP_{it}	-12.11***	-10.81***	-9.38***	-9.21***
ΔURB_{it}	-7.47***	-2.59***	-2.13**	-4.36***

Notes: Table entries are the t_b^* statistic of Moon and Perron (2004) computed for a given number of factors $m = 1, 2, 3, 4$. Under the null, the t_b^* statistic tends to a standard normal distribution as $T, N \to \infty$ and $N/T \to 0$.
The one-sided 10%, 5%, and 1% critical values are -1.282, -1.645, and -2.327, respectively. '*', '**', and '***' denote significance at the 10%, 5%, and 1% levels, respectively.

Table 7: Pooled mean group estimates of the adjustment coefficients and long-run parameters, 1975-2005

Country	ϕ		RDI		RLIR		POP		URB		R^2_{adj}	Half-life, years
BE	-0.079	**	1.052	***	-0.028	***	0.484	***	3.223	***	0.596	8.4
	(0.045)		(0.129)		(0.004)		(0.103)		(1.022)			
CA	-0.197	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.766	3.2
	(0.035)		(0.129)		(0.004)		(0.103)		(1.022)			
FI	-0.116	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.564	5.6
	(0.041)		(0.129)		(0.004)		(0.103)		(1.022)			
FR	-0.276	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.792	2.1
	(0.048)		(0.129)		(0.004)		(0.103)		(1.022)			
DE	-0.031	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.720	22.0
	(0.010)		(0.129)		(0.004)		(0.103)		(1.022)			
IE	-0.043	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.758	15.8
	(0.014)		(0.129)		(0.004)		(0.103)		(1.022)			
IT	-0.175	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.255	3.6
	(0.050)		(0.129)		(0.004)		(0.103)		(1.022)			
JA	-0.012		1.052	***	-0.028	***	0.484	***	3.223	***	0.365	57.4
	(0.053)		(0.129)		(0.004)		(0.103)		(1.022)			
NL	-0.204	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.619	3.0
	(0.047)		(0.129)		(0.004)		(0.103)		(1.022)			
ES	-0.108	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.567	6.1
	(0.037)		(0.129)		(0.004)		(0.103)		(1.022)			
SE	-0.091	**	1.052	***	-0.028	***	0.484	***	3.223	***	0.704	7.3
	(0.039)		(0.129)		(0.004)		(0.103)		(1.022)			
CH	-0.005		1.052	***	-0.028	***	0.484	***	3.223	***	0.410	138.3
	(0.016)		(0.129)		(0.004)		(0.103)		(1.022)			
UK	-0.158	***	1.052	***	-0.028	***	0.484	***	3.223	***	0.839	4.0
	(0.036)		(0.129)		(0.004)		(0.103)		(1.022)			
US	-0.026	*	1.052	***	-0.028	***	0.484	***	3.223	***	0.671	26.3
	(0.017)		(0.129)		(0.004)		(0.103)		(1.022)			

Note: Figures in parentheses are the standard errors.

Figure 1: Classification of countries according to the real house price growth (house price indices divided by the CPI, 100=1975), 1975-2004

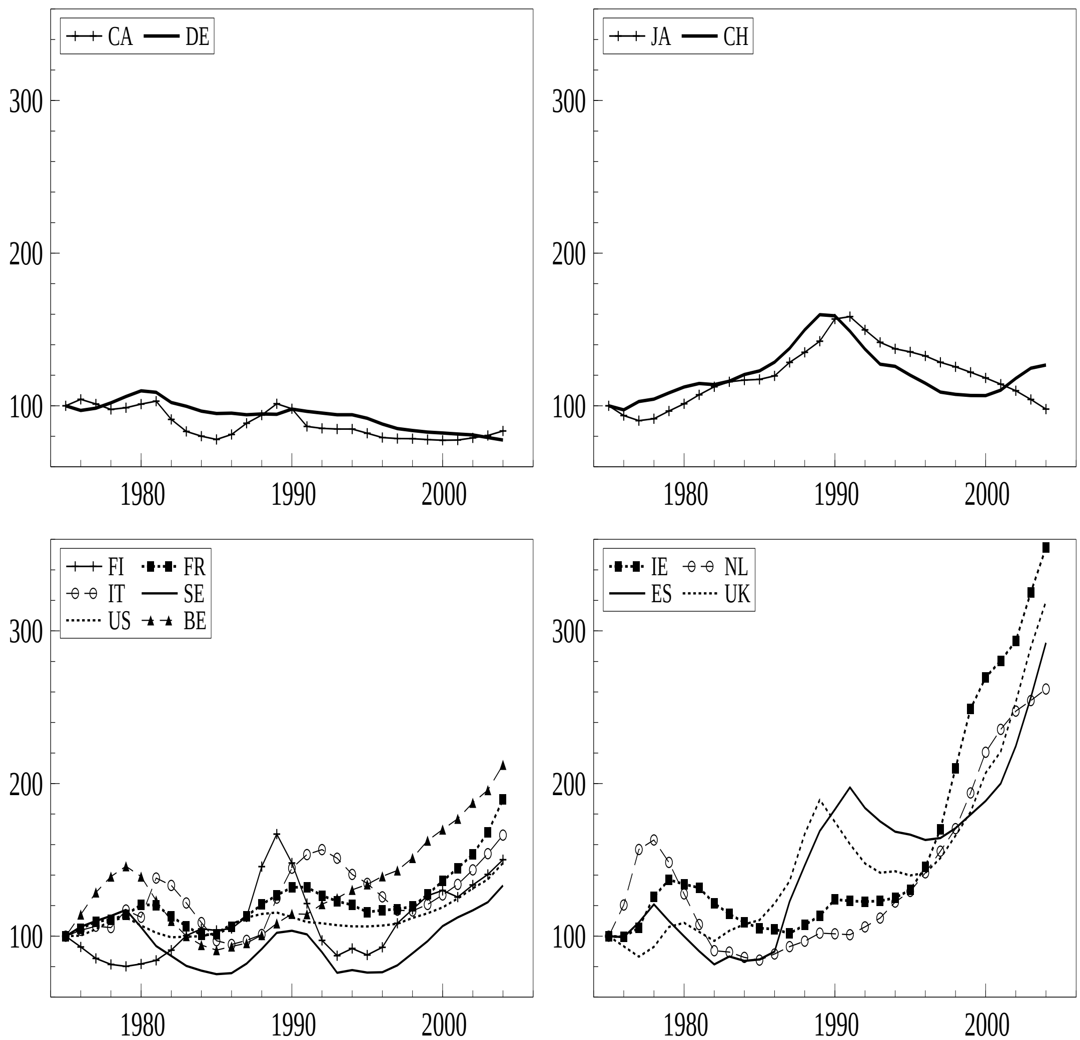

Figure 2: Stability check of adjustment parameters, ϕ, 1975-2004

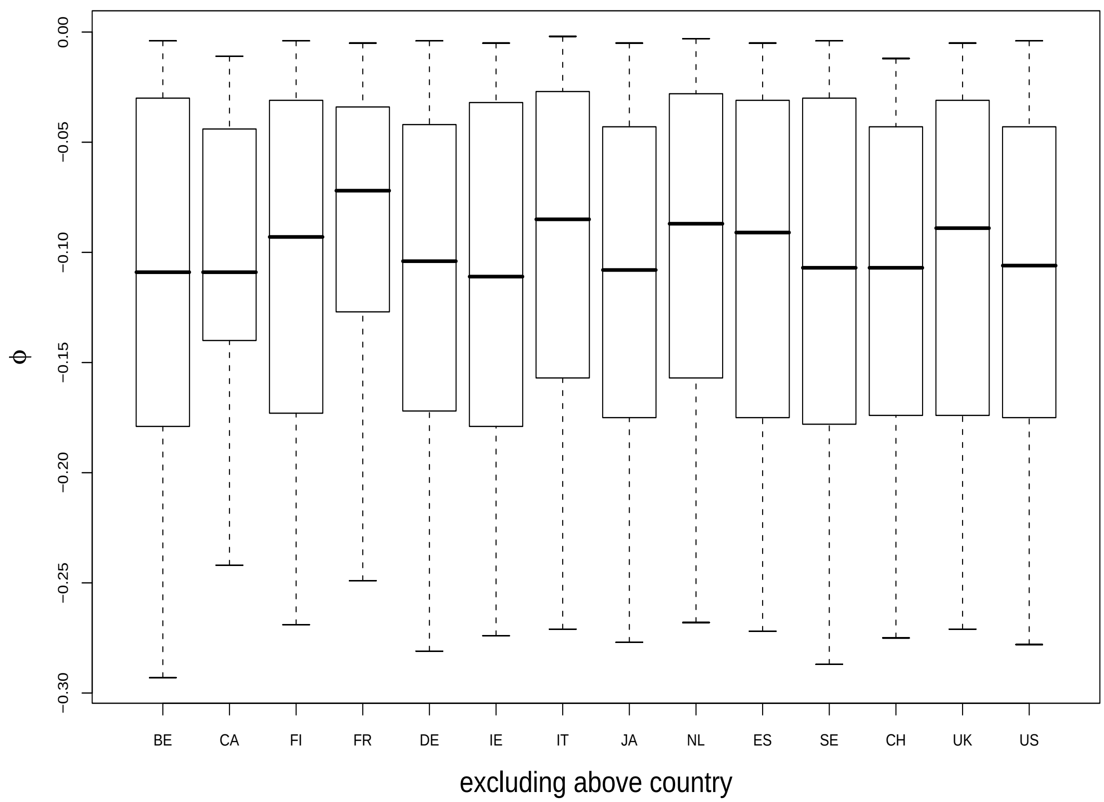

Figure 3: Stability check of long-run parameters, 1975-2004

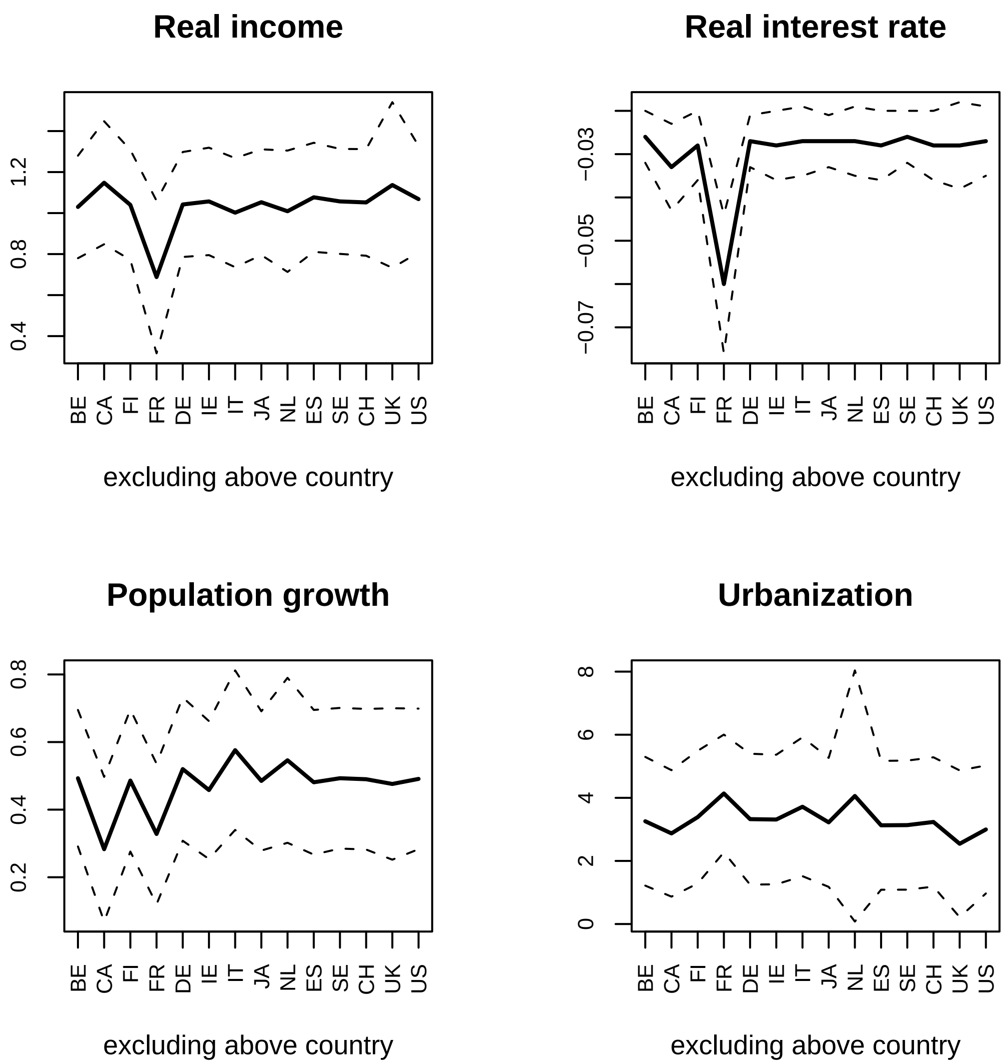

Figure 4: Housing market fundamentals in 14 selected OECD countries, 1975-2004