European Commission

ISSN 2443-8022 (online)

Regional Spillovers in the Hungarian Housing Market: Evidence from a Spatio-Temporal Model

Gábor Márk Pellényi

DISCUSSION PAPER 095 | APRIL 2019

EUROPEAN ECONOMY

Economic and Financial Affairs

European Economy Discussion Papers are written by the staff of the European Commission's Directorate-General for Economic and Financial Affairs, or by experts working in association with them, to inform discussion on economic policy and to stimulate debate.

The views expressed in this document are solely those of the author(s) and do not necessarily represent the official views of the European Commission.

Authorised for publication by Manfred Bergmann, Director for Economies of the Member States I.

This paper exists in English only and can be downloaded from https://ec.europa.eu/info/publications/economic-and-financial-affairs-publications_en.

Luxembourg: Publications Office of the European Union, 2019

PDF ISBN 978-92-79-77432-4 ISSN 2443-8022 doi:10.2765/420868 KC-BD-18-022-EN-N

European Commission
Directorate-General for Economic and Financial Affairs

Regional Spillovers in the Hungarian Housing Market

Evidence from a Spatio-Temporal Model

Gábor Márk Pellényi

Abstract

This paper analyses housing market spillovers between Hungarian regions in a small, spatio-temporal model, which features both the house price and housing supply as endogenous variables. The paper estimates both the long-run relationship between housing variables and economic fundamentals, and the short-run adjustment path of the housing market towards the long-run equilibrium. Long-run elasticities are in line with previous studies. The size of spillovers between Hungarian regions is economically meaningful; therefore, region-specific developments such as the recent run-up of Budapest house prices can have significant aggregate effects.

JEL Classification: E32, R12, R21, R31.

Keywords: housing demand, housing supply, overvaluation, regional spillover, ripple effect.

Acknowledgements: I thank Jean-Charles Bricongne, Violaine Faubert, Norbert Gaál and Dino Pinelli for their useful comments and suggestions.

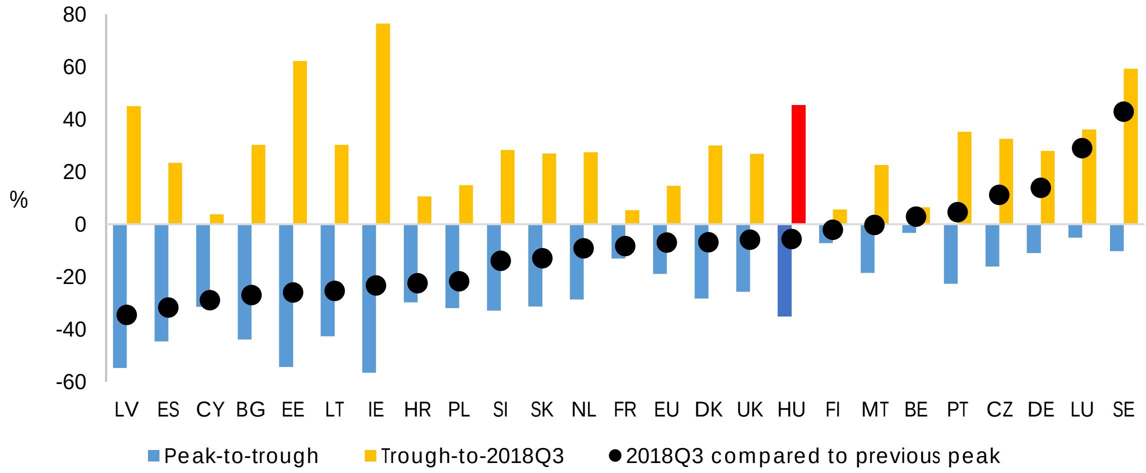
Contact: Gábor Márk Pellényi, European Commission, Directorate-General for Economic and Financial Affairs, gabor-mark.pellenyi@ec.europa.eu.

CONTENTS

1. INTRODUCTION

The Hungarian housing market saw large swings since the crisis. House prices, adjusted for changes in purchasing power, fell by 35% from their pre-crisis peak until 2013. This decline significantly exceeded the average 19% peak-to-trough decrease in the EU (see Graph 1). The recovery of real house prices has also been strong in European comparison. Since bottoming out, the national, real house price index soared by 45% until the third quarter of 2018, already raising some concerns about overvaluation. Residential construction underwent similar changes: the number of housing units built fell from 35-40 thousand per year in the pre-crisis period to just 7.3 thousand in 2013. Construction has recovered somewhat slower than prices: the number of new housing units rose to 14.4 thousand in 2017. Nonetheless, the number of new building permits has already risen back to the range of 35-40 thousand.

Graph 1. **Real house price changes over the last housing cycle in EU Member States**



Note: nominal house price index for all housing units, deflated with the harmonised consumer price index.

Source: Eurostat

Behind strong aggregate figures lie great regional differences within Hungary. The capital, Budapest is well ahead of the rest of the country, both in terms of price increases and housing construction. Thus, worries about potential overvaluation mainly concern Budapest (MNB, 2018). It remains a question whether overvaluation in one region can have significant spillovers for the rest of the country. International evidence suggests that it can; such a 'ripple effect' has been identified first in the UK, then in several countries (see Meen, 2016 and its citations, as well as Oikarinen, 2004 for Finland; Hviid, 2017 for Denmark; Teye et al., 2017 for the Netherlands, among others). Regional spillovers of house prices are of interest for several reasons including regional economic convergence or divergence, interregional mobility and financial stability. The last issue is relevant in several European countries where house price increases in capital cities have decoupled from the rest of the countries (Claeys et al., 2017).

The long-term evolution of residential construction is another important question. In recent years, construction activity in Hungary has been boosted by the cyclical recovery of income and credit, and government measures (including reduced value added tax rate on housing until 2019 and the expansion of housing subsidies to families). However, due to long project cycles, property developers already need to have a clearer view of housing demand over the longer run. Similarly, banks need to be aware of the longer-term outlook to maintain prudent lending standards. Demographic trends are expected to have a very strong impact on the long-term demand for housing, but these trends can be very uneven across regions.

Thus, regional heterogeneity and spillovers are important for both house prices and construction. However, only few papers deal simultaneously with these issues. Existing studies that jointly analyse prices and quantities on the housing market typically use aggregate data; e.g. Meen (2000), Gattini and Hiebert (2010). Papers explicitly modelling spatial linkages generally focus on the behaviour of prices and treat housing stock as exogenous; e.g. Beenstock and Felsenstein (2010). A joint analysis of regional house prices and volumes appears in the spatial VAR model of Beenstock and Felsenstein (2007) but contrary to this paper, they ignore the possibility of cointegration, which can make their estimates of long-run effects less reliable. Housing starts are also present in the spatial VAR of Brady (2014) but his impulse response analysis only concerns house prices. The present paper goes beyond these studies by quantifying the effect of region-specific house price shocks on aggregate construction activity.

This paper also adds to the literature on the Hungarian housing market. Earlier studies focused on aggregate developments, and were preoccupied with estimating key elasticities of house prices; see e.g. Kiss and Vadas (2007), Horváth (2008), Berki and Szendrei (2017). The supply side of the market, its regional dimension and the role of demography have received less attention. The VECM model of Berki and Szendrei (2017) incorporated both house prices and the housing stock but their analysis was restricted to the behaviour of prices. Regional studies include Farkas (2011) who estimated the effect of demographic change on house prices in a detailed regional panel. Székely (2008) and Békés et al. (2016) examined the causes of house price differentials between cities. However, these regional studies did not consider spillovers. Therefore, this paper also contains novel results for Hungary, especially concerning the elasticity of housing supply and regional linkages.

Finally, the paper complements existing spatial econometric analyses by considering an economic rationale for spatial linkages. Researchers have so far typically assumed simple spatial patterns such as the physical neighbourhood of territorial units. However, this approach has been criticised for lacking economic underpinning; see Corrado and Fingleton (2012). This paper applies a novel approach to construct the spatial matrix, by making use of the matrix of interregional migration flows, published by the Hungarian Central Statistical Office.

The estimated parameters of model are similar to the results of earlier, aggregate studies. By taking into account the sluggish reaction of supply, the model reproduces the intuition that housing market cycles are quite long and price fluctuations are large. According to the model, house prices have returned to levels consistent with economic fundamentals by 2017. Budapest appears overvalued by a significant margin, at least partly due to the rising activity of foreign investors in the capital city. This Budapest-specific overvaluation can boost construction activity throughout the country: spillovers can stimulate as much construction activity outside Budapest as inside the capital. Such large effects warrant attention, especially because the behaviour and the motivations of foreign real estate investors are poorly understood. Badarinza and Ramadorai (2018) show that foreign property investors transmit macroeconomic shocks in their host countries to investment destinations, making Budapest potentially more sensitive to changes in international risk factors.

The regional analysis of house price valuations can complement the standard toolkit of macroeconomic surveillance under the aegis of the Macroeconomic Imbalance Procedure of the EU. It can help detect localised bubbles which do not yet appear in the aggregate, national house price index. A model featuring regional spillovers can also help determine whether local house price imbalances have significant aggregate consequences, which may endanger financial and macroeconomic stability. Finally, the spatial analysis of the housing market may also inform cohesion policy which aims to moderate regional differences in the level of economic development.

2. EMPIRICAL APPROACH

The analysis is based on a small, dynamic model of regional housing markets, explicitly accounting for temporal and regional relationships, and describing the behaviour of both prices and quantities. This section lays out a simple stock-flow model with geographic spillovers, starting from basic economic theory.

The equation of house prices can be derived from standard models of housing demand, see e.g. Muellbauer and Murphy (2008). The demand for housing services per household (h) is determined by real income (y), the real price of housing (p) and other demand shifters (z; all variables are in logs):

$$h = -\alpha p + \beta y + z \qquad (1)$$

In this framework, housing is a consumption good, which can be substituted with other consumption items.[1] Their relative demand depends on their relative price, which is the nominal house price index divided by the price level of private consumption, i.e. the real house price index. Muellbauer and Murphy (2008) also argue that since housing is a durable good, its demand depends on its user cost. They suggest a definition of user cost (uc) that takes into account depreciation (δ), financing costs (r), taxes (t), and expected revaluation ($\dot{p}^e$) capturing expected capital gains: $uc = \delta + r + t - \dot{p}^e$. Let z_1 denote the vector of demand shifters different from user cost. As housing supply is inflexible in the short run, equation (1) is usually rearranged for prices and treated as an inverted demand curve:

$$p = \frac{1}{\alpha}(\beta y - h - \theta \cdot (\delta + r + t - \dot{p}^e) + z_1) \qquad (2)$$

Note that if the above equation characterises a rational expectations (or “steady state” or “long run”) equilibrium, then expected house price changes must be zero. Furthermore, the amortisation rate can be assumed constant. Then, if property taxes are low, or if they do not change much, financing costs alone may be sufficient to characterise user cost. Clearly, several papers have found that house price expectations are not rational but backward-looking (see Muellbauer, 2012 for a survey). Still, this issue can be addressed in a dynamic formulation of the house price equation, where the parameters of lagged variables will pick up backward-looking expectations of house price changes. For this reason, the baseline estimation will use only the cost of mortgage finance, and the user cost will appear in a robustness check.

Other demand shifters may include demography (e.g. the size of younger, first-home buyer cohorts) and credit availability. Furthermore, in a spatial setting, the house prices of neighbouring regions also affect housing demand through the location choice of households; see Beenstock and Felsenstein (2010) for a formal derivation. Adding subscripts, we arrive at the following long-run equation:

$$\log p_{i,t} = \frac{1}{\alpha}\left(\beta \log y_{i,t} - \log s_{i,t} + \gamma \log pop_{i,t} - \theta \cdot r_t + \rho \log p^f_{i,t}\right) \qquad (3)$$

Where $p_{i,t}$ is the real house price index in region i in year t; $y_{i,t}$ is real income per capita, $s_{i,t}$ is the housing stock, $pop_{i,t}$ is population, r_t is the cost of mortgage finance, and $p^f_{i,t}$ is the weighted average real house price level of neighbouring regions.

In the short-run house prices may deviate significantly from this equilibrium, but a tendency to revert to the equilibrium can be expected. In econometric terms it means that house prices should be cointegrated with the above fundamentals, and to follow an equilibrium correction process.

[1] While housing is often considered a basic necessity, some amenities can still be substituted with purchased goods and services (e.g. guest rooms with hotels, a washing machine with laundry services, even a kitchen with restaurant meals).

The derivation of equation (3) assumed that the housing stock is exogenous. In the long run this is clearly not the case. We therefore need another equation to describe housing supply. The modelling of housing supply is less straightforward than the derivation of the demand curve. The literature has put forward two main approaches. In the asset pricing model of Poterba (1984) housing investment is determined by the *level* of house prices relative to construction cost, which in turn reflects the net present value of future housing services. On the other hand, the urban growth model of Mayer and Somerville (2000) posits that construction is a function of house price *changes*, not the price level. They argue that housing comprises of two elements, structure and land. Structure is elastically supplied, but land is inelastic even in the long run because the amount of land inside any city is limited. Therefore an exogenous increase in housing demand will lead to a permanently higher house price. The jump to this higher price level will encourage developers to convert available land to residential use. Mayer and Somerville (2000) argue that this approach is empirically more valid because the house price level is typically nonstationary, while the series of housing starts is stationary.

Empirically the two approaches can be reconciled (Grimes and Aitken, 2010). If house prices and the housing stock are cointegrated, then net construction (the change in housing stock) may be driven by the deviation of house prices from the cointegrating relationship, which is a stationary variable. In an equilibrium correction framework net construction can also be affected by current and lagged house price changes. Other factors including income and demography may also influence changes in the housing stock, even in the short run.

At the end, we arrive at a bivariate equilibrium-correction model which can flexibly accommodate several theories of the housing market:

$$\begin{bmatrix}\Delta p_{i,t}\\ \Delta s_{i,t}\end{bmatrix} = -\begin{bmatrix}\gamma_1\\ \gamma_2\end{bmatrix}\left(p_{i,t-1} - \beta_s s_{i,t-1} - \boldsymbol{\beta}'_X \mathbf{X}_{i,t-1}\right) + \begin{bmatrix}0 & \omega_s\\ \omega_p & 0\end{bmatrix}\begin{bmatrix}\Delta p_{i,t}\\ \Delta s_{i,t}\end{bmatrix} + \boldsymbol{\Omega}_1\begin{bmatrix}\Delta p_{i,t-1}\\ \Delta s_{i,t-1}\end{bmatrix} + \cdots +$$
$$+\boldsymbol{\Omega}_q\begin{bmatrix}\Delta p_{i,t-q}\\ \Delta s_{i,t-q}\end{bmatrix} + \mathbf{A}_0\Delta\mathbf{X}_{i,t} + \cdots + \mathbf{A}_r\Delta\mathbf{X}_{i,t-r} + \begin{bmatrix}\epsilon_{i,t}\\ \eta_{i,t}\end{bmatrix} \tag{4}$$

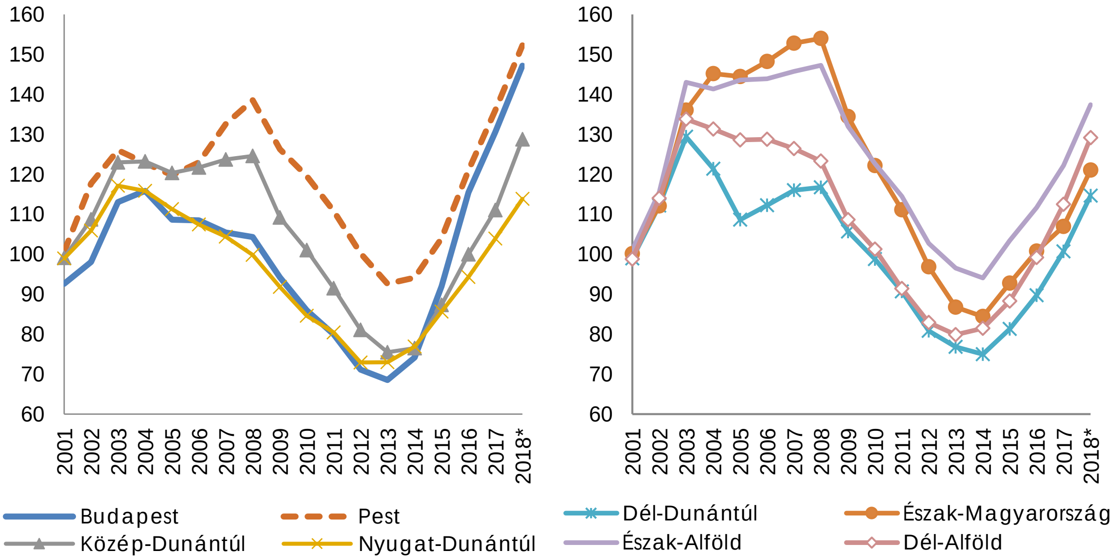
where $\mathbf{X}_{i,t} = (p^f_{i,t}, r_t, y_{i,t}, pop_{i,t})$. Neighbourhood effects are captured by a spatial matrix ($\mathbf{W}$) which links each region to its neighbours through region-specific weights: $\boldsymbol{p}^f_t = \mathbf{W}\boldsymbol{p}_t$. Equation (4) is a general form of the model. Its exact specification and estimation will be presented in Section 4. Note that while the model captures the main macroeconomic drivers of housing markets, it omits some relevant factors, notably household indebtedness, government subsidies and construction costs. These are all very relevant for Hungary (see e.g. Verner and Gyöngyösi, 2018, for debt or Vadas, 2009, for subsidies), but cannot be easily incorporated in the model due to the lack of reliable, regional data.

3. DATA

Quality-adjusted house price data (denoted with p) are compiled by Magyar Nemzeti Bank for the seven NUTS2 regions; the Central Hungary region is also divided into Budapest (the capital) and the surrounding Pest County which contains most of the suburbs of Budapest (Graph 12 in the Annex contains a map of the country). Quarterly data are available since 2001. These house price indexes are based on actual transactions as reported in the land registry; they are cleaned from changes in the composition of transactions, and are deflated with consumer prices (see Graph 2).

The spatial matrix **W** plays an important role because it determines the strength of linkages between individual regions. The spatial econometric literature often uses a simple contiguity matrix which assigns zeros to non-neighbouring regions and ones to neighbours; the matrix is then row-normalised. Contiguity matrices are simple but they may lack economic interpretation (see Corrado and Fingleton, 2012, for an overview of the criticism regarding the choice of the spatial matrix).

Graph 2. **Real house prices by region (2001Q1=100)**



Note: adjusted for changing composition of housing transactions, deflated with consumer prices. * estimation based on data until the third quarter of 2018.

Source: Magyar Nemzeti Bank

Instead of using the contiguity matrix, this paper introduces a spatial matrix based on interregional migration flows, which could better capture the economic linkages between regional housing markets. In Hungary, 95% of housing units are occupied by their owner, according to the 2011 census. If a person wants to move from region A to B, she must typically sell her property in region A and acquire a new home in region B. If a region-specific house price shock increases prices in A, the purchasing power of the migrating person will increase in B; this wealth effect will bid up prices in B. The strength of this wealth effect can be assumed proportional to the number of people changing residence from A to B. Thus, the w_{ij} element of the spatial matrix represents the number of people moving from region j to region i as a percentage of the population of region i; these flows are averaged over the years 2001-2016 (see Table 1). This method is similar to the approach of Holly et al. (2011) who make use of commuting data when constructing the spatial matrix.

Table 1. **Spatial matrix based on interregional migration flows**

Region	Budapest	Pest	Közép-Dunántúl	Nyugat-Dunántúl	Dél-Dunántúl	Észak-Magyaro.	Észak-Alföld	Dél-Alföld
Budapest	0.0000	0.3327	0.1194	0.0683	0.0801	0.1468	0.1487	0.1040
Pest	0.6397	0.0000	0.0641	0.0242	0.0332	0.0838	0.0885	0.0664
Közép-Dunántúl	0.3105	0.1395	0.0000	0.1741	0.1220	0.0764	0.0852	0.0923
Nyugat-Dunántúl	0.2128	0.0726	0.2821	0.0000	0.1526	0.0885	0.0995	0.0918
Dél-Dunántúl	0.2894	0.0986	0.1968	0.1548	0.0000	0.0531	0.0607	0.1466
Észak-Magyarország	0.3651	0.1818	0.0715	0.0472	0.0322	0.0000	0.2390	0.0632
Észak-Alföld	0.3142	0.1710	0.0691	0.0472	0.0350	0.2259	0.0000	0.1376
Dél-Alföld	0.2833	0.1765	0.1062	0.0588	0.1117	0.0752	0.1883	0.0000

Note: The average gross inflow to each row region between 2001 and 2016 is broken down by source region (in columns), and normalised to one. For example, the value 0.3327 in the second column of the first row means that 33.3% of the gross inflow to Budapest originated from Pest county.

Source: Hungarian Central Statistical Office

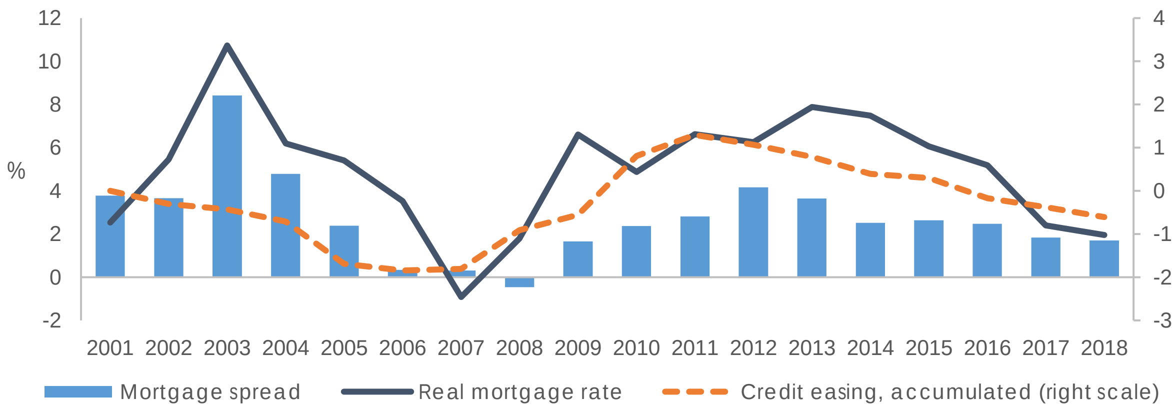
Real income per capita (y) is net real wage income per head in the 15-64 years age group. Average net wages are reported by the Central Statistical Office for each region; these are multiplied by regional employment and divided by population aged 15-64 years, and by the consumer price index. The main reason for using wages instead of total income is the lack of regional breakdown for the latter before 2010. Focusing on labour income may still give a reliable picture, because it accounts for 60-75% of total income in every region, and it is responsible for the bulk of income variation between regions. A robustness check will also use regional GDP as a proxy for income.

Data for the 15-64 years old population (pop) and the housing stock (s) are available at annual frequency at the Central Statistical Office. The series contain structural breaks around the 2011 census. These level shifts were corrected by applying the average growth rate of population and housing stock in the neighbouring years for the year when the break occurs.

The choice of the demographic variable may be contested by arguing that housing decisions are taken at the household level, not at the individual level. Therefore the number of households may be a more appropriate measure. Indeed, the two variables show different trends: while population fell 6% between 2001 and 2015, the number of households rose by 4.1%. However, there are several arguments against using household data. A theoretical reason is that the number of households is endogenous because household formation reacts to housing market developments; see e.g. Haurin et al. (1993). Therefore the relevant decision unit for housing choice may still be the individual. A practical reason is that regional data on the number of households is available only in five-year intervals (in census or micro-census years). Simple interpolation between these years may be misleading because household formation can also react significantly to business cycles in the intervening years (see Lee and Painter, 2013, for evidence).

Credit conditions are measured by the spread of the mortgage rate over 10-year government bonds ($spread$, see Graph 3). The mortgage rate is the weighted average lending rate of local and foreign currency housing loans, including all charges above interest, weighted by monthly contracted volumes. Local currency rates also take into account the effect of a government interest subsidy scheme, which played a particularly important role in the early 2000s.

Graph 3. **Mortgage lending conditions in Hungary**



Note: Real mortgage rate = rate on new loans in HUF, EUR or CHF including all charges, weighted by new contracted volumes – current-year CPI inflation. Mortgage spread = nominal mortgage rate – 10-year government bond yield. Credit easing, accumulated = self-reported changes in banks' mortgage lending conditions, cumulated since 2000, similar to the approach of Aron et al. (2012).

Source: Magyar Nemzeti Bank, author's calculations

The mortgage spread can be a more useful measure than the interest rate, for two reasons. First, it is more comprehensive because it better reflects banks' willingness to provide credit. Second, it is less affected by measurement error. Lending and borrowing decisions should be based on the ex ante real interest rate, which requires the measurement of inflation expectations. Hungarian inflation was volatile during the sample period, affected first by a disinflationary trend then by several one-off price level shocks (e.g. indirect tax changes). Therefore, expected inflation cannot be calculated easily and the resulting real interest rate can be very noisy. In contrast, the calculation of the mortgage spread requires only nominal interest rates. Robustness check will examine alternative measures of credit cost in Section 5.

All data except for spreads are in logs; mortgage spreads are in percentage points. The dataset covers the years 2001 to 2017, with eight cross-sectional units. Although the time span appears short, it roughly covers one full housing cycle; in addition, the peaks and troughs are heterogeneous across regions. Thus, the sample may still be sufficient to estimate key elasticities.

4. ESTIMATION RESULTS

4.1. LONG-RUN RELATIONSHIPS

The first step of the analysis is the estimation of equation (3). In order to motivate the regional panel framework, it should be checked whether house prices and explanatory variables are indeed correlated across regions. The bias-corrected Lagrange Multiplier test of Baltagi et al. (2012) has the null hypothesis of no cross-sectional correlation. The CD test of Pesaran (2015) has the null of no or weak (non-pervasive) cross-sectional correlation, against an alternative hypothesis of strong (pervasive) cross-sectional dependence (see Chudik et al. 2011 for exact definitions of weak and strong dependence). The null hypothesis is rejected for both tests (see Table 2), which points to strong cross-sectional dependence between regional variables. Bailey et al. (2016) argue that spatial econometric models can only account for weak cross-sectional dependence; the data should first be purged of strong dependence to avoid misleading inference. This can be done by extracting common factors from regional variables, or by adding observed variables which can account for strong dependence (such as income or credit cost, which are driven largely by economy-wide shocks). This paper takes the latter approach by explicitly modelling the determinants of regional house prices, while also including spillover effects from neighbouring regions.

Table 2. **Tests of cross-section dependence**

	Level		Differenced	
Variable	Bias-corr. LM	CD	Bias-corr. LM	CD
p	43.44 ***	18.73 ***	42.48 ***	18.62
y	58.27 ***	21.59 ***	48.01 ***	19.72 ***
pop	38.22 ***	9.09 ***	30.99 ***	12.06 ***
s	59.25 ***	21.75 ***	44.24 ***	18.96 ***

Note: null hypothesis of no cross-sectional dependence (LM test) or at most weak cross-sectional dependence (CD test) rejected at *10%, **5%, ***1% significance level.

In order to estimate the long run equation, the time series properties of the data also need to be tested. Standard panel unit root tests may be biased under cross-sectional correlation; this problem can be overcome by using the cross-sectionally augmented CIPS test of Pesaran (2007) for all variables except mortgage spreads. This test augments standard (A)DF regressions for each cross-section unit with lags and differences of the cross-sectional averages of the tested variable. The CIPS test statistic is the average of the t-statistics for the lagged explanatory variable across cross-section units. Pesaran

(2007) reports critical values and finds that the test performs adequately even in small samples such as the one in this paper. According to the CIPS test all variables can be considered I(1) (see Table 3).

Table 3. **Unit root tests for model variables**

Variable	Test type	Level	Differenced
p	CIPS	-2.06	-2.27 *
y	CIPS	-2.47 **	-2.73 ***
pop	CIPS	-1.83	-2.33 *
s	CIPS	-1.69	-2.89 ***
spread	ADF	-1.96	-3.49 **

Note: null hypothesis of unit root rejected at *10%, **5%, ***1% significance level. Critical values of the CIPS test are from Pesaran (2007), for the case N=10, T=15; regressions include a constant, no trend.

After the preliminary analysis of the variables, the cointegrating relationship can be tested and estimated (see Table 4). The literature has proposed various techniques. Beenstock and Felsenstein (2010) show that the OLS estimator remains superconsistent when temporal and spatial cointegrations are both present. Kao and Chiang (2000) recommend using dynamic OLS because of its smaller bias in small samples. Gonzalo (1994), among others, recommends an autoregressive distributed lag (ADL) specification over two-step estimators of error correction models. However, ADL models are sensitive to the appropriate specification of short-term dynamics, which may be problematic in my sample with just 17 time series observations. Finally, SUR models (both in levels and in the ADL form) can also take into account any residual (weak) cross-sectional dependence that may affect parameter estimates.

Table 4. **Estimates of the long-run relationship**

	OLS	DOLS	ADL	SUR	ADL SUR
p^f	0.774 ***	0.511 ***	0.104	0.801 ***	0.342 ***
	(0.057)	(0.079)	(0.214)	(0.013)	(0.097)
y	0.229 ***	0.388 ***	0.861 ***	0.211 ***	0.635 ***
	(0.062)	(0.114)	(0.268)	(0.016)	(0.133)
pop	1.865 ***	2.472 ***	4.680 ***	1.831 ***	4.057 ***
	(0.333)	(0.301)	(0.910)	(0.065)	(0.414)
s	-1.149 ***	-2.241 ***	-4.801 **	-1.038 ***	-3.334 ***
	(0.322)	(0.494)	(1.484)	(0.089)	(0.604)
spread	-0.530 *	-1.241 ***	-3.183 *	-0.254 ***	-1.779 **
	(0.283)	(0.349)	(1.160)	(0.079)	(0.467)
Total effects					
y	1.01	0.79	0.96	1.06	0.96
pop	8.25	5.06	5.23	9.22	6.17
s	-5.08	-4.58	-5.36	-5.23	-5.07
spread	-2.34	-2.54	-3.55	-1.28	-2.70
Adjusted R^2	0.888	0.940	0.884	0.989	0.914
Number of observations	136	120	120	136	120
Cross-sectional dependence tests of residuals					
Bias-corrected LM test	9.35	6.53	1.04	-3.16 **	-2.88 ***
CD test	-0.69 ***	2.29 *	1.45	-1.70 *	-1.17
Unit root test of residuals					
CIPS test	-2.23 ***	-1.83 ***	-1.40	-2.31 ***	-1.50

Note: Dependent variable is *p*. All models include regional fixed effects. Parameters are significant / null hypothesis rejected at *10%, **5%, ***1% significance level. Critical values of the CIPS test with the null hypothesis of a unit root are from Pesaran (2007), for the case N=10, T=15; regressions include no deterministics. Calculated total effects take into account regional spillovers through the spatial weight matrix.

Although individual parameter estimates vary somewhat, they paint a consistent picture regarding the main determinants of regional house prices. Real income, population, the housing stock and mortgage spreads all have significant coefficients; their signs are consistent with the predictions of theory. Aggregate estimates typically find that the income elasticity of house prices is close to one. Note that with spatial spillovers the total effect of income is not simply β_y, but $N^{-1}\boldsymbol{\iota}'\left(\mathbf{I}-\beta_{p_f}\mathbf{W}\right)^{-1}\mathbf{I}\beta_y\boldsymbol{\iota}$, where $\mathbf{I}$ is an $N \times N$ identity matrix, and $\boldsymbol{\iota}$ is an $N \times 1$ vector of ones. Point estimates of the total effect are presented in the second block of Table 3. Indeed, including all interregional spillovers, the income elasticity of house prices is very close to unity.

The impact of mortgage spreads is also economically meaningful, and is in line with earlier estimates for Hungary by Kiss and Vadas (2007). Over the last credit cycle, mortgage spreads changed 400 bps both in the expansion and contraction phases; this magnitude is associated with roughly 10% change in real house prices according to the SUR specification. Finally, the effect of demography and housing supply on prices is very strong. Although estimated coefficients for population are high, they are not without precedent. In a much-disputed paper, Mankiw and Weil (1989) presented similar elasticities for demography. Furthermore, while Takáts (2012) estimated the elasticity of house prices to total population near one in a cross-country panel, the parameters for individual countries were often as high as 10. As for the housing stock, Berki and Szendrei (2017) estimate an elasticity of 3, qualitatively similar to the current estimate of 5. Thus, while the estimates in this paper should be taken with a grain of salt, they are consistent with earlier evidence. These results are also robust to the exclusion of the large recession in 2009 following the financial crisis, or the omission of all years after 2010 which could be affected by the revaluation of foreign currency debt and the resulting debt overhang (see also Section 5 and Table 5), although the exact parameter values obviously change somewhat.

Statistical tests hint that the explanatory variables removed strong cross-sectional dependence from regional house price data, but some weak dependence may remain. The null hypothesis of the CD test (no or at most weak dependence) is either rejected or is close to rejection in all models. However, null of any (weak or strong) dependence cannot be rejected according to the bias-corrected LM test.

The cointegrating nature of the estimated long-run relationships is checked by testing the stationarity of the residuals. Again, given the apparent weak cross-sectional dependence across residuals, the CIPS panel unit root test is recommended. It suggests that the residuals from the OLS, DOLS and SUR estimations are stationary. In the case of the ADL specifications the null of unit root cannot be rejected, although by a narrow margin.

4.2. DEVIATIONS FROM THE LONG-RUN EQUILIBRIUM

The estimated long-run relationships can be used to calculate house price levels which are consistent with the actual values of those economic fundamentals which are included in the model. If all regions are in equilibrium, house prices are given by:

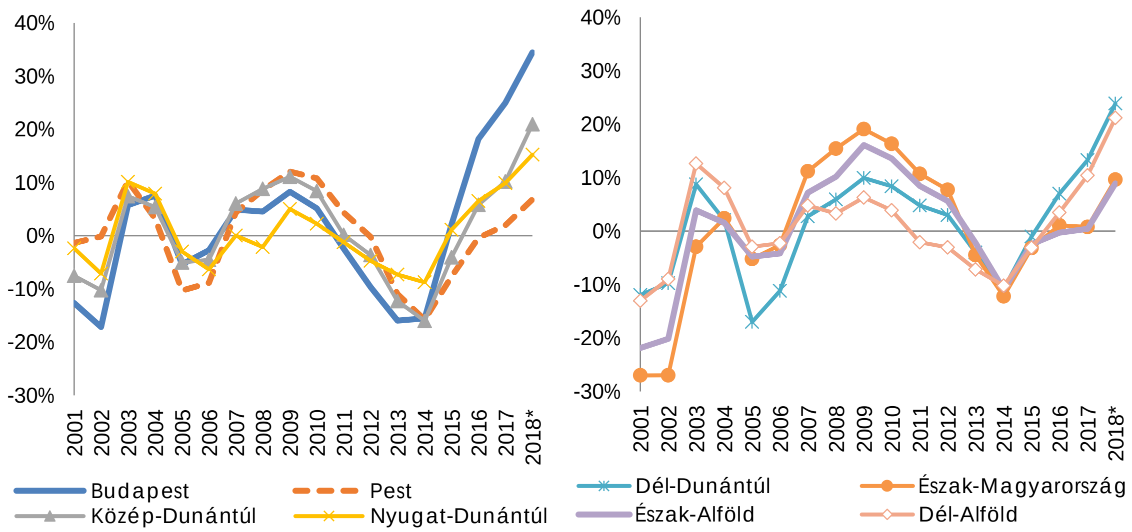
$$\boldsymbol{p}_t^* = \mathbf{X}_t\boldsymbol{\beta} + \beta_{p_f}\mathbf{W}\boldsymbol{p}_t^* \tag{5}$$

Where matrix $\mathbf{X}_t$ contains all explanatory variables for each region, except for the weighted average house price of neighbouring regions ($\boldsymbol{p}_t^*$), and vector $\boldsymbol{\beta}$ contains the parameters associated with these variables. Rearranging, we get

$$\boldsymbol{p}_t^* = \left(\mathbf{I}-\beta_{p_f}\mathbf{W}\right)^{-1}\mathbf{X}_t\boldsymbol{\beta} \tag{6}$$

Graph 4 illustrates the deviation of regional house prices from their fundamental value, calculated with equation (6). To guard against misspecification, the deviations for all five estimated equations have been calculated, and their averages are presented. As the one-step estimation of the full error correction model does not separately identify the region-specific constant of the cointegrating relationship, deviations from the equilibrium in each region are normalised to average zero in the 2001-2017 period. Preliminary estimates for 2018 are also shown, based on real house prices until the third quarter of 2018.

Graph 4. **House price deviations from fundamentals**



Note: Average of five estimation methods presented in Table 4. * estimation based on data until the third quarter of 2018.

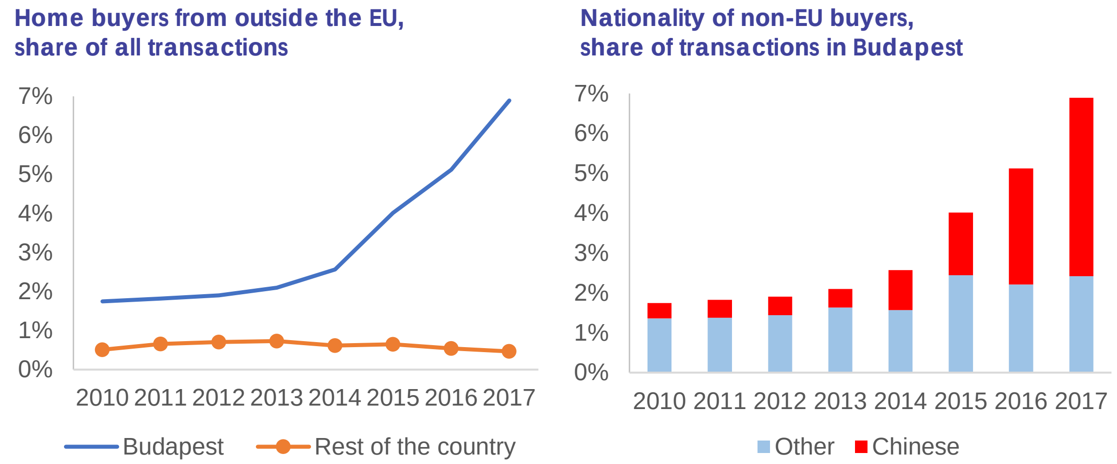
The 2000s started with undervalued house prices. Rising income levels, a generous interest subsidy scheme and later a foreign-currency lending boom contributed to rapid price increases before the crisis (Vadas, 2009). House prices overshot fundamentals by 10-20% by 2008. A gradual correction followed the crisis; by 2014, house prices were undervalued by around 10% in most regions. Since then, prices recovered, and by 2018 they already appear to exceed values that are consistent with standard macroeconomic fundamentals. The most striking case is Budapest, where prices are now 35% above their fundamental value. The possible overvaluation of the Budapest market is also indicated by the fact that Budapest prices have risen by 130% from their post-crisis trough until the third quarter of 2018, well above the national average of 73%. A model excluding Budapest would not change the above conclusions for the remaining regions.

Similar valuation gaps are prepared regularly within the European Commission (based on the model of Philipponet and Turrini, 2017); and by Magyar Nemzeti Bank in its Housing Market Report (which uses the method of Berki and Szendrei, 2017). This paper uses the same fundamentals as the model-based methods of the European Commission and MNB, although the exact choice of variables differs somewhat. MNB (2018) also publishes results for Budapest, which is derived from a model containing Budapest only. In contrast, the model proposed in this paper allows the joint estimation of valuation gaps in all regions. Its results correspond with the findings of Philipponnet and Turrini (2017) that Hungarian house prices were undervalued after the financial crisis, but became aligned with fundamentals around 2015. MNB (2018) estimates a higher fundamental price level, resulting in persistent undervaluation for the national economy. Still, it also finds that Budapest has reached much higher valuation levels than the rest of the country. Based on the analysis of house price levels, Bricongne et al. (2019) find that overvaluation can be established when the price of a 100 m^2 apartment exceeds ten years' worth of household income. An average-sized apartment in Budapest is much smaller, just 67 m^2 according to the 2016 microcensus. Still, even for this apartment size, the price-to income ratio rose from 5.6 in 2013 to 10.5 in 2017, signalling risk of overvaluation.

Why have prices exceeded fundamentals by 2018 in most regions? There are three possible explanations. First, government subsidies for home purchases were expanded significantly in 2015 and added to housing demand. The value of newly issued, subsidised loans doubled between 2015 and 2017. Second, capacity constraints in the construction sector increased construction costs, which may have filtered into existing home prices. Finally, in 2018 it became clear that the government would not extend the temporary value added tax reduction on new housing, which could have contributed to recent price increases.[2] Unfortunately, these factors could not be included in the model because no reliable regional data is available on the take-up of housing subsidies or homebuilding costs.

What explains the recent, extraordinary growth of Budapest prices? One possibility is that the Budapest market is more sensitive to income and interest rates than the rest of the country. The capital is a major economic, touristic and educational hub, where the rental market is relatively more developed. Therefore, Budapest may be a particularly attractive location for property investors. This paper does not analyse parameter heterogeneity across regions due to the small sample size. Another possibility is that some omitted factors are responsible for Budapest-specific price increase. Some suggestive evidence for the role of investors comes from data on house purchases by extra-EU buyers. These buyers must obtain permission from local authorities; data on their requests show rising interest in the Budapest market, especially on behalf of Chinese individuals. By 2017 extra-EU buyers accounted for almost 7% of residential real estate purchases in Budapest, with Chinese buyers playing a prominent role (see Graph 5). This coincided with increasing capital outflows from China, which affected house prices in several other countries as well. The special visa programme of the Hungarian government between 2013 and 2017 may have also contributed to rising extra-EU interest.

Graph 5. **Residential real estate purchases by non-EU citizens**



Note: the graphs are based publicly available registration data of extra-EU individuals. They are divided by the number of housing market transactions in the respective region.

Source: Government of Hungary, Hungarian Central Statistical Office

Data on intra-EU buyers are scarce because these individuals do not need to seek permits since 2006, while the Central Statistical Office only started systematic data collection on foreign buyers in 2017 based on land registry data. These recent CSO figures can be compared with the data collections of Illés and Michalkó (2008) for 2001-2006. It turns out that the number of foreign buyers in Budapest in

[2] The standard value added tax rate of 27% was reduced to 5% for new housing built between 1 January 2016 and 31 December 2019. Market participants were uncertain until 2018 whether the temporary tax cut would be made permanent. With the adoption of tax laws for 2019 it became clear that the tax cut is indeed temporary. Nevertheless, in order to ease capacity shortages, the tax cut was extended until the end of 2023 for those projects that had already received a building permit by November 2018.

2017 more than doubled compared to the average between 2001 and 2006; their share of all transactions increased even more. On the contrary, outside Budapest, the number of foreign buyers in 2017 was still somewhat below the 2001-2006 average. Graph 5 suggests that this divergence between Budapest and the rest of the economy was largely driven by extra-EU demand.

Another, partly overlapping development is the increasing use of housing for non-residential purposes in Budapest. Jancsik et al (2018) found that the number of accommodations offered through Airbnb tripled between 2014 and 2017; by the end of 2017 housing units used for flatsharing accounted for 1% of all apartments in Budapest. According to their research, half of these apartments belonged to real estate investors.

4.3. THE ESTIMATED MODEL

This section describes the estimation of the full housing market model in equation (4). To limit the number of parameters some parameters were set to zero based on economic reasoning; variables with insignificant parameters were also dropped. The final specification is the following (*ecm* refers to the deviation from the cointegrating vector):

$$\Delta p_{i,t} = f\left(ecm_{i,t-1}, \Delta y_{i,t}, \Delta spread_{t-1}, \Delta p_{i,t-1}\right)$$

$$\Delta s_{i,t} = g\left(\Delta s_{i,t-1}, \Delta y_{i,t}, \Delta pop_{i,t}, \Delta p_{i,t-1}\right) \tag{7}$$

The short-run dynamics of house prices are determined by a correction towards the long-run equilibrium, as well as short-term changes in the fundamentals. This specification assumes that only income and mortgage spreads affect short-run dynamics. Population and the housing stock may be considered slow-moving variables, which affect house prices with a longer lag, and do not necessary influence its short-term fluctuations. Finally, lagged house price changes capture possibly adaptive expectation of agents and any other sources of price stickiness.

In the second ("supply") equation of the model, housing stock is driven by income, population and lagged house prices, while showing some inertia due to the time lags of construction projects. In the case of housing stock, error correction towards the long-run equilibrium was not significant, confirming the intuition that supply is sticky and house prices play the key role in market clearing. However, the role of lagged house price changes in housing supply dynamics (i.e. construction activity) was significant. The estimated model is therefore reminiscent of the theoretical approach of Mayer and Somerville (2000).

There are no simultaneous links between the two equations, because both include only predetermined variables. Therefore the equations can be estimated separately. In theory, endogeneity may arise in the house price equation because the explanatory variables include current income. Two separate arguments can address this concern. First, this equation has an autoregressive distributed lag specification (where cross-sectional units are grouped together), which yields consistent OLS estimates even with endogenous regressors, if residuals are not autocorrelated (Pesaran and Shin, 1998). Indeed, standard tests do not find serial correlation in the residuals of the estimated model. Second, housing wealth effects are estimated to be very small in Hungary (see Kiss and Vadas, 2007), thus the contemporaneous link from house prices through consumption to income is weak. In the case of the housing supply equation, the small share of construction in GDP should ensure that changes in building activity have a very limited effect on current-year income, which may again alleviate concerns of reverse causality.[3] The robustness of parameter estimates was also checked using maximum likelihood (not reported), but the choice of estimation method did not have a material effect.

The parameters of the cointegrating relationship are taken from the SUR specification in Table 4. To account for remaining cross-section dependence, the dynamic equations are also estimated with SUR. The estimated model takes the following form (standard errors are in parentheses below the parameter values; fixed effects in the price equation are not reported):

$$\Delta p_{i,t} = \underset{(0.039)}{-0.313}\left(p_{i,t-1} - \underset{(0.097)}{0.342}\, p^{f}_{i,t-1} - \underset{(0.133)}{0.635}\, y_{i,t-1} - \underset{(0.604)}{3.333}\, pop_{i,t-1} + \underset{(0.467)}{1.779}\, s_{i,t-1} + \underset{(0.467)}{1.780}\, spread_{t-1} \right) + \underset{(0.063)}{0.521}\, \Delta y_{i,t} - \underset{(0.288)}{0.492}\, \Delta spread_t + \underset{(0.053)}{0.553}\, \Delta p_{i,t-1} + \mu_i + \epsilon_{i,t}$$

$adj.R^2 = 0.644$

$$\Delta s_{i,t} = \underset{(0.032)}{0.772}\, \Delta s_{i,t-1} + \underset{(0.001)}{0.004}\, \Delta y_{i,t} + \underset{(0.016)}{0.128}\, \Delta pop_{i,t} + \underset{(0.001)}{0.005}\, \Delta p_{i,t-1} + \underset{(0.000)}{0.001} + \eta_{i,t}$$

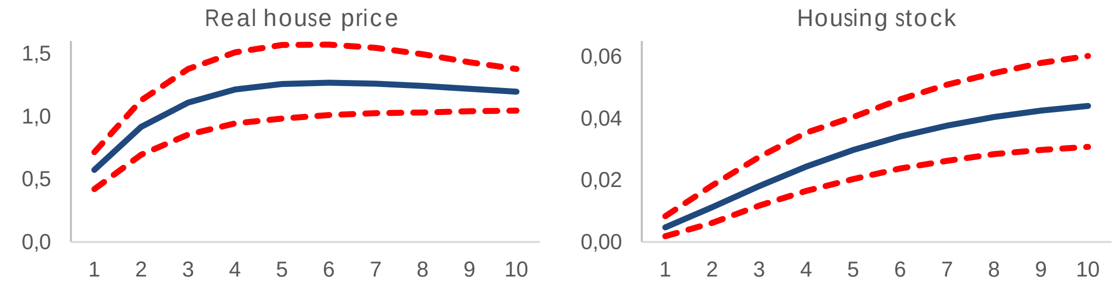
$adj.R^2 = 0.898$

4.4. DYNAMIC SIMULATIONS

Dynamic simulations can help to better understand the behaviour of the system, and check whether the estimated parameters are intuitive. Impulse responses are calculated to permanent shocks of the main exogenous variables (income, population, mortgage spreads), and to a temporary residual shock of the house price equation.

The responses to the shocks of exogenous variables are shown on Graphs 6-8. Impulse responses are the same for each region because these shocks are identical in every region, and estimated model parameters do not differ across regions. The response of house prices to income shocks is somewhat hump-shaped, peaking around year 5. Then, as housing supply begins to adjust, prices moderate. A 1% increase in real income lifts real house prices by around 1% in the long run, and it raises the housing stock by 0.04% (see Graph 6). For illustration, a steady 3% annual income growth would result in a 5000 unit net increase (construction less demolition) of dwellings every year. To put this number in perspective, on average 24 thousand dwellings were built per year over the sample period.

Graph 6. **The response of house prices and housing stock to a permanent 1% rise of real income (%)**

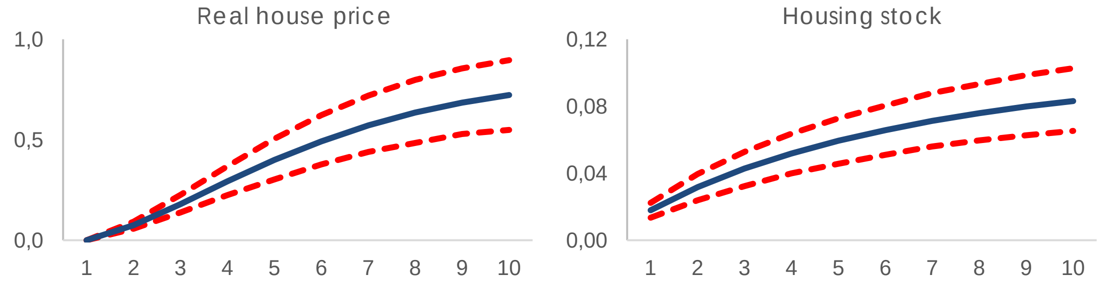


Note: horizontal axis denotes time in years, vertical axis is measured in percentages. Mean response (blue solid line) and 95% confidence intervals (red dashed line), based on 1000 bootstrap repetitions of the estimated model.

[3] Total construction activity, including all input linkages, is estimated to have accounted for 8% of gross value added in 2017. However, only 16.4% of construction output was related to dwellings in 2017, thus the share of housing production in value added could have been around 1-1.5%.

If working age population (15-64 years) grew by 1%, house prices would rise by 0.7% in the long term, while the housing stock would grow by 0.8% (see Graph 7). Over the last decade, the working age population of Hungary decreased by an annual average of 0.5%. Such a demographic trend would be consistent with an annual net decline of 18 thousand dwellings. Thus, given recent economic and demographic trends, the model suggests that the current housing stock in Hungary is sufficient, and there may be little need for a net increase in housing. Three caveats should be added to this result. First, the model does not account for the amortisation and the quality of housing, which may require continuous investments even with decreasing population. Nonetheless, such investments do not require a net change in housing stock (i.e. the number of existing housing units). Second, the model is silent about possible asymmetries in housing supply. For example, houses are being built when population grows, but they are not automatically demolished when population decreases. The analysis of asymmetries is beyond the scope of this paper due to small sample size. The third caveat is that the determinants of household formation are poorly understood. If population decline is coupled with an autonomous decrease in household size (for example due to changing preferences in society), the model may underestimate housing demand.

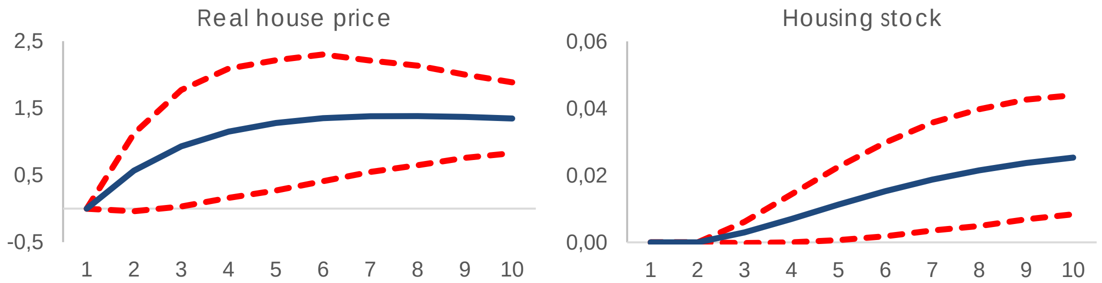
Graph 7. **The response of house prices and housing stock to a 1% population increase (%)**



Note: horizontal axis denotes time in years, vertical axis is measured in percentages. Mean response (blue solid line) and 95% confidence intervals (red dashed line), based on 1000 bootstrap repetitions of the estimated model.

Mortgage spreads are ultimately linked to credit cycles, which are relatively long-lived. Financial market regulations (i.e. stronger competition among banks) may even cause permanent changes in spreads. A permanent (or highly persistent), 100bp decrease in mortgage spreads would raise real house prices by approximately 1.4% in the long run, and it would boost the housing stock by 0.03% (see Graph 8). During recent credit cycles in the Hungarian economy, mortgage spreads moved by 400 bps. Such a decrease in spreads would boost real house prices by over 5%, while the net impact on the housing stock may reach 4000 units. This is not negligible compared to the 37 thousand unit change in construction from the peak to the trough of the last housing cycle.

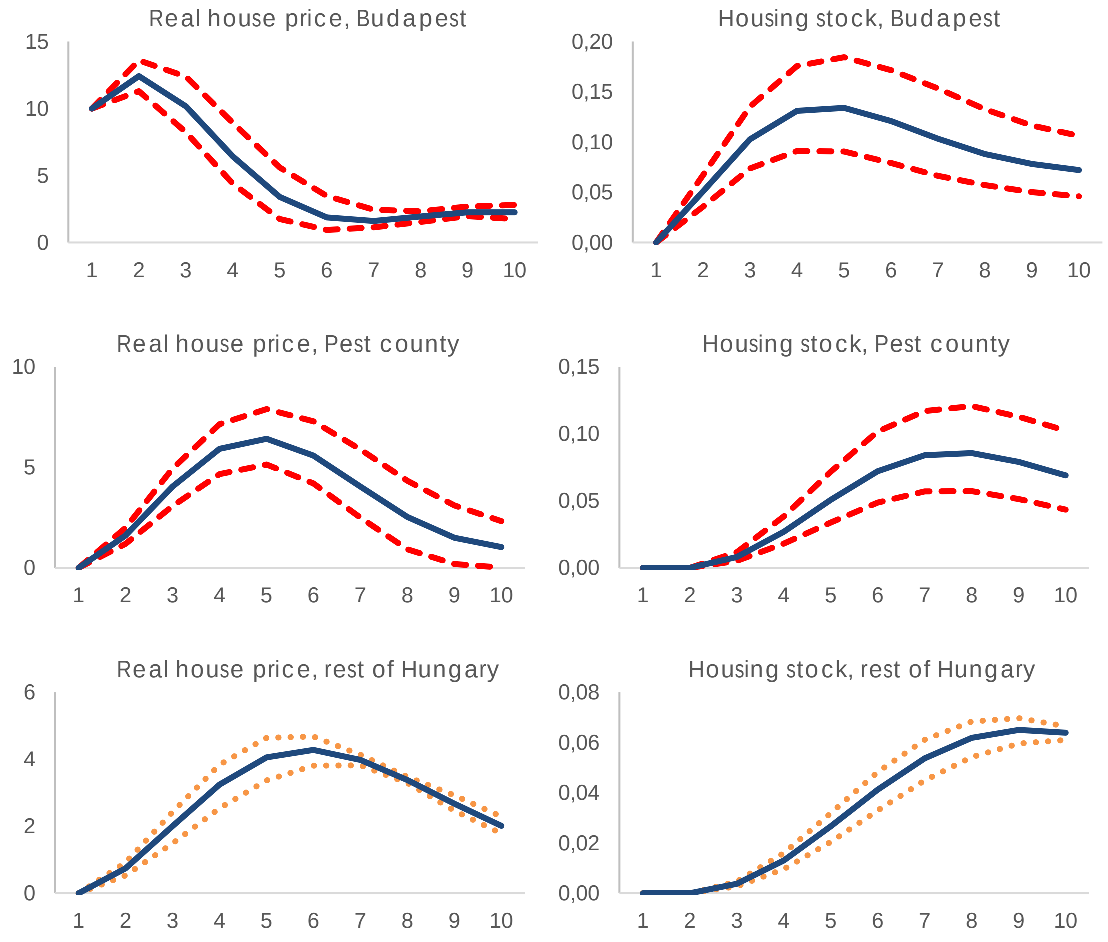
Graph 8. **The response of house prices and housing stock to a permanent 100 bp decrease of the mortgage spread (%)**



Note: horizontal axis denotes time in years; vertical axis is measured in percentages. Mean response (blue solid line) and 95% confidence intervals (red dashed line), based on 1000 bootstrap repetitions of the estimated model.

Finally, Graph 9 shows the impact of a Budapest-specific house price shock. This shock is of particular interest because it demonstrates the spatial mechanics of the model. It is also of policy relevance given the potential overheating of the Budapest real estate market. To save space, 95% confidence intervals are plotted only for Budapest and Pest County; for the remaining regions only the minima, averages and maxima of mean responses across regions are shown. A temporary, exogenous rise in Budapest prices takes five years to dissipate. It transmits to all regions of Hungary; the effect is the strongest in neighbouring Pest County. The temporary price increase also boosts housing construction across the country; the housing stock rises by over 0.1% in Budapest and by 0.08% in Pest County.[4] Altogether, the housing stock would increase by approx. 1200 units in Budapest by the fourth year, and by over 500 in the rest of the country. The response outside Budapest appears to keep rising in the later years, to 2300 units by the eighth year.

Graph 9. **The response of house prices and housing stock to a temporary 10% increase of Budapest house prices (%)**



Note: horizontal axis denotes time in years; vertical axis is measured in percentages. Mean response (blue solid line) and 95% confidence intervals (red dashed line), based on 1000 bootstrap repetitions of the estimated model. Responses for the rest of Hungary are averages for the remaining six regions; orange dotted lines show the minimum and maximum responses across these regions.

The significant response of prices outside Pest County may come as a surprise. One could argue that an omitted common factor could explain the co-movement of prices in Budapest and more distant

[4] The long-run decrease of the housing stock could be due to the physical amortisation of existing dwellings. However, it can also point to asymmetric responses to positive and negative shocks.

regions. Tests of cross-sectional dependence show no evidence of such omitted factors.[5] Such "long-distance" spillovers may be explained by second-round effects from Pest County to its neighbours, and also by direct migration from Budapest to more distant regions. Given its large size and central role, Budapest is a dominant source of migration for each Hungarian region (see Table 1).

Recall that one possible reason for the recent fast appreciation of Budapest housing is the inflow of foreign investment. Inasmuch as this explanation is true, the model suggests that foreign buyers in Budapest can influence prices and building activity in the entire country. This is in line with the results of Liao et al. (2015), who find that real estate investments by foreigners in Singapore have significant ripple effects in neighbourhoods where the share of foreign buyers was lower.

As a caveat, note two results that might be artefacts of the model. First, the effect of the Budapest-specific shock on prices is quite long-lived, contrary to earlier results for different countries (e.g. Hviid, 2017, for Denmark). This is due to the strong and persistent reaction of prices in the rest of the country. The 'ripple effect' is reflected back to Budapest, keeping prices high even after the original shock has faded. In the long term however, the response of prices could still revert to zero. Second, while Budapest and neighbouring Pest country show marked differences in responses, the remaining regions react very similarly to a Budapest-specific shock. This is a consequence of identical parameters and only limited differences in regional weights. A robustness check of Section 5 will relax some of these restrictions.

5. ROBUSTNESS CHECKS

The first group of robustness tests concerns the choice of explanatory variables (see Table 5). The SUR specification of the long-run relationship is reestimated after changing some of the variables. The first potential concern is that population and housing stock data show structural breaks after 2010. In the baseline estimation these level shifts were corrected manually, which may cause measurement error. Thus, the cointegrating equation was reestimated with data up to only 2010; parameters did not change materially. The income variable could be challenged because net wage income per capita omits potentially important income elements (social benefits, entrepreneurial income) which may show different dynamics than wages. A more comprehensive measure of income, per capita GDP at the regional level is also tried, deflated with consumer prices. Again, the signs and of parameters remain unaffected, but the (total) elasticity of house prices to income is smaller than one, which is the typical result of aggregate studies.

Several alternatives to mortgage spreads can be proposed. Of these, the level of real mortgage rates does not yield significant parameters. Lending conditions are calculated similar to Aron et al. (2012), by cumulating changes in banks' self-reported mortgage lending conditions since 2000, based on the quarterly lending survey of the central bank. This measure has three advantages over mortgage spreads. It takes into account non-price lending conditions; it is less affected by changes in borrower quality (which can bias the average level of mortgage spread); and it may better reflect macroprudential regulations which mainly intervene in non-price conditions. However, it also has the disadvantage that changes in lending conditions are measured by the market shares of banks reporting changing standards, and not by the actual amount of easing or tightening. This can create large measurement errors. In this specification, the main parameters of interest remain unchanged, with the exception of the housing stock. However, the significance of lending conditions themselves is somewhat weaker than is the case for mortgage spreads.

[5] Pesaran's CD test statistic for the residuals of the dynamic house price equation is 0.167, which means that the null hypothesis of at most weak cross-section dependence cannot be rejected in favour of an alternative hypothesis of strong dependence. Therefore this test does not indicate a missing common factor.

The next financial variable is a standard measure of user cost which takes the average house price change of the last four years as a proxy for future price expectations, based on the suggestion of several papers surveyed by Muellbauer (2012). As regional house prices are only available from 2001, lagged national house price changes are used in the first four years of the sample. The calculation of the user cost should ideally take into account property-related taxes and benefits, which in Hungary include the stamp duty; real estate taxes which only apply to certain properties (e.g. holiday homes) and only in certain municipalities; and the deductibility of mortgage payments from personal income tax until 2006. The first two are ignored in this analysis because stamp duties were broadly stable throughout the sample, and there is no publicly available information on municipal property taxes. Regarding the tax advantage of borrowing, the Statistical Office publishes annual data on the average size of newly issued mortgages; this value is assumed to be the same in every region and furthermore, homebuyers are assumed to fully utilise this tax benefit. For simplicity, the depreciation rate for housing is set to 2%. As the above description already suggests, there are several potential sources of measurement error in the calculation of user cost, which reduce its empirical appeal. Indeed, this model gives counterintuitively weak effects for income and the housing stock.

Table 5. **Robustness checks of the long-run relationship**

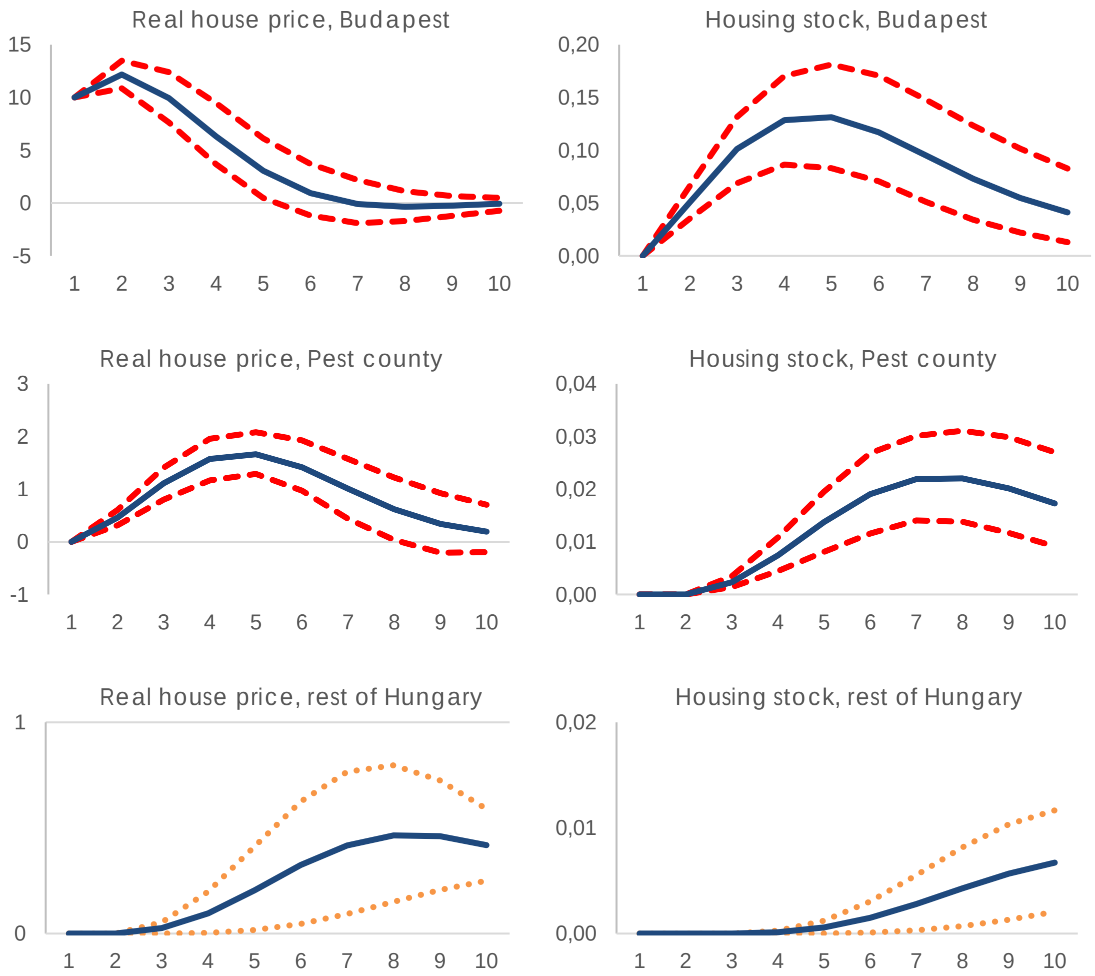
	Baseline	Until 2010	GDP	Real rate	Lending conditions	User cost	$W_{contiguity}$
pf	0.801 ***	0.716 ***	0.872 ***	0.815 ***	0.807 ***	0.789 ***	0.808 ***
	(0.013)	(0.008)	(0.009)	(0.011)	(0.010)	(0.011)	(0.017)
y	0.211 ***	0.364 ***		0.196 ***	0.183 ***	0.068 ***	0.321 ***
	(0.016)	(0.010)		(0.013)	(0.010)	(0.023)	(0.017)
gdp			0.068 ***				
			(0.014)				
pop	1.831 ***	1.225 ***	1.354 ***	1.769 ***	1.723 ***	1.348 ***	1.842 ***
	(0.065)	(0.098)	(0.054)	(0.055)	(0.051)	(0.079)	(0.064)
s	-1.038 ***	-0.991 ***	-0.364 ***	-0.877 ***	-0.800 ***	-0.086	-1.310 ***
	(0.089)	(0.087)	(0.041)	(0.066)	(0.052)	(0.128)	(0.098)
spread	-0.254 ***	-0.271 ***	-0.121 ***				-0.351 ***
	(0.079)	(0.015)	(0.040)				(0.071)
real mortgage rate				-0.066			
				(0.043)			
lending conditions					-0.004 **		
					(0.002)		
user cost						-0.243 ***	
						(0.042)	
Total effects							
y / gdp	1.06	1.28	0.53	1.06	0.95	0.32	1.67
pop	9.22	4.31	10.58	9.56	8.93	6.39	9.59
s	-5.23	-3.49	-2.84	-4.74	-4.15	-0.41	-6.82
Lending variable	-1.28	-0.95	-0.95	-0.36	-0.02	-1.15	-1.83
Adjusted R^2	0.989	0.999	0.997	0.992	0.996	0.994	0.996
Number of observations	136	80	128	136	136	136	136
Cross-sectional dependence tests of residuals							
Bias-corrected LM test	-3.16 **	-3.968 ***	-3.705 ***	-3.261	-3.788 ***	-3.835 ***	-3.441 ***
CD test	-1.70 *	-0.450	-0.518	-1.207	-0.576	-0.390	-0.360
Unit root test of residuals							
CIPS test	-2.31 ***	-1.210	-2.165 ***	-2.561 ***	-2.526 ***	-2.702 ***	-1.681 *

Note: Dependent variable is *p*. All models include regional fixed effects. Parameters are significant / null hypothesis rejected at *10%, **5%, ***1% significance level. Critical values of the CIPS test with the null hypothesis of a unit root are from Pesaran (2007), for the case N=10, T=15; regressions include no deterministics. Calculated total effects take into account regional spillovers through the spatial weight matrix.

Unit root tests confirm the stationarity of residuals in the robustness checks above, except – not surprisingly – for the shorter sample. In all, the first batch of robustness tests confirms that the long-run relationship between house prices and fundamentals exists, and that spillovers from neighbouring regions are significant, regardless of the exact choice of variables.

The next robustness check concerns the **W** matrix of spatial connectivity. This is the only source of regional heterogeneity in impulse responses because model parameters are assumed constant across regions due to data limitations. The baseline choice for the spatial matrix is based on interregional migration which can be a proxy for wealth spillovers through the selling and buying of property in different regions. It remains an imperfect measure of spillovers because not all migrating individuals sell and buy residences (e.g. university students moving to a dormitory in Budapest then back home). Meanwhile, income spillovers through commuters are not captured by this particular weighting scheme. It is therefore important to understand the sensitivity of results to the choice of matrix **W**. A natural candidate is the contiguity matrix, a standard starting point in spatial econometrics, where neighbouring regions are encoded with ones and non-neighbours with zeroes, and row sums are then normalised to unity.

Graph 10. **Robustness check: the response of house prices and housing stock to a temporary 10% rise of Budapest house prices with a contiguity matrix (%)**



Note: horizontal axis denotes time in years; vertical axis is measured in percentages. Mean response (blue solid line) and 95% confidence intervals (red dashed line), based on 1000 bootstrap repetitions of the estimated model. Responses for the rest of Hungary are averages for the remaining six regions; orange dotted lines show the minimum and maximum responses across these regions.

Results for the long-run relationship with the contiguity matrix are similar to the baseline, and neighbourhood effects remain significant. The estimated total elasticities (including spillovers) are stronger than in the baseline model (last column of Table 5). The full model was also re-estimated to to check the sensitivity of spillovers from Budapest-specific price shocks to the choice of the spatial matrix (see Graph 10). The overall pattern remains similar to the baseline model, although the size of spillovers is somewhat smaller. In the baseline, the reaction of Pest county is about half of the response of Budapest (both for prices and quantities), while the average reaction in the remaining regions is about one-third. When **W** is a contiguity matrix, the reaction of Pest county is one-fifth of the response in Budapest, while the reaction in remaining regions is less than one-tenth (albeit still significant). With sufficiently long time series it would be possible to test the significance of **W** matrix elements which were set a priori, or even estimate them. This approach was taken for example by Bailey et al. (2016), although their time series were about 8 times as long as in this paper. Thus, until longer series become available for Hungary, only the existence of spatial spillovers appears robust, but their exact magnitude can be estimated with some uncertainty.

The constancy of parameters across regions is another strong assumption of the model. Again, small sample size does not allow a detailed analysis of parameter heterogeneity, but it can be checked whether excluding one region at a time changes results for the long-run relationship. For brevity and easier comparison, only the calculated total effect of each explanatory variable is reported, taking into account spillovers (see Table 6). It appears that house prices in Budapest and Pest County might be somewhat less sensitive to local income, but they could be more affected by mortgage spreads, than prices across the rest of the country. This is consistent with the larger role of speculative housing demand in Budapest, and it might also suggest that credit conditions play an important role in suburban sprawl around Budapest.

Table 6. **Robustness checks of the long-run relationship, excluding individual regions**

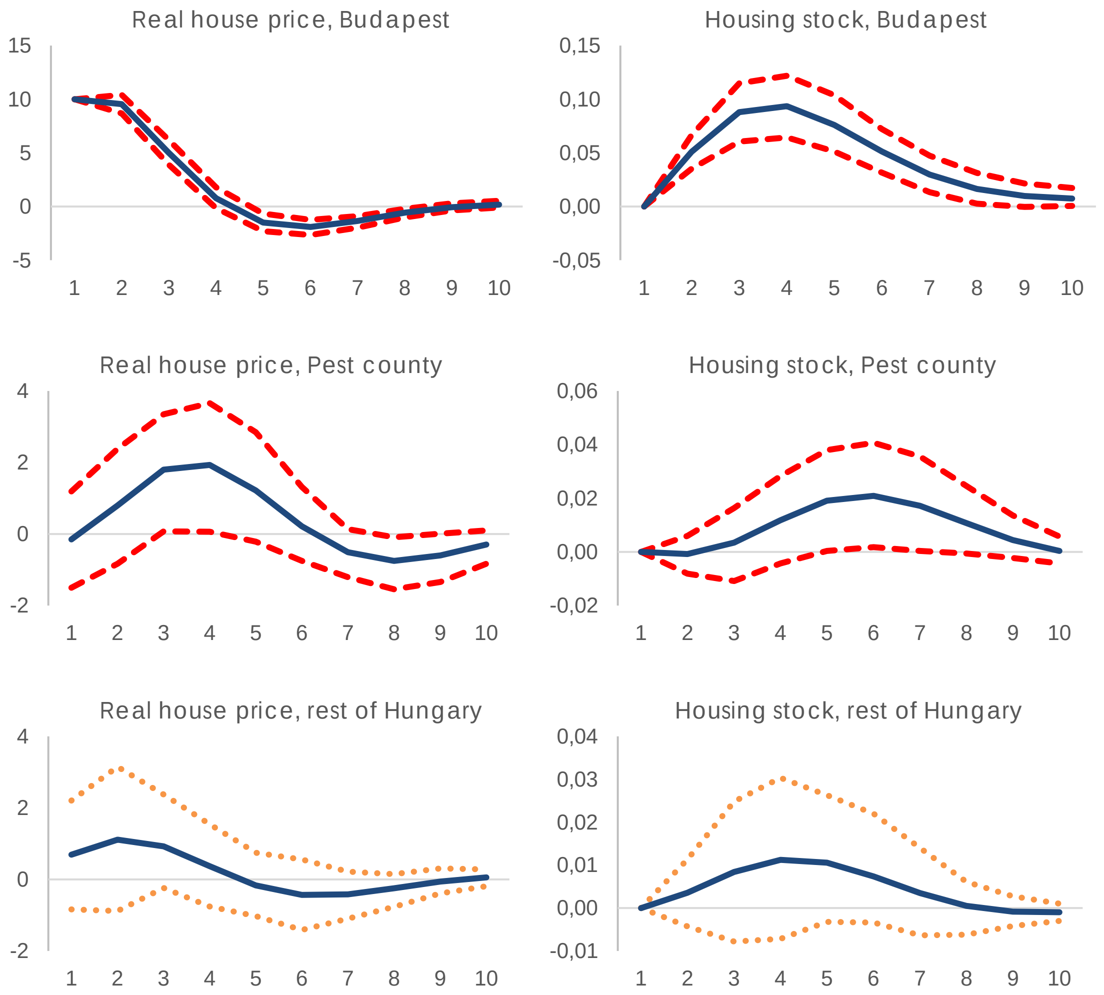
	Total effect excluding region								Total effect with all regions
	Budapest	Pest	Közép-Dunántúl	Nyugat-Dunántúl	Dél-Dunántúl	Észak-Magyaro.	Észak-Alföld	Dél-Alföld	
y	1.34	1.19	0.91	1.01	1.01	0.93	0.90	0.85	1.06
pop	8.52	9.88	7.72	10.45	8.54	8.42	7.82	7.89	9.22
s	-4.36	-1.61	-4.82	-4.34	-5.16	-5.10	-4.93	-4.71	-5.23
spread	-0.64	-0.29	-1.02	-1.44	-0.84	-0.73	-0.93	-1.19	-1.28

Note: The estimated model is the SUR specification of the cointegrating vector (column 4 in Table 4). In each column, the highlighted region was excluded from the sample, and the spatial matrix was rescaled for the remaining regions. Calculated total effects take into account regional spillovers through the spatial weight matrix.

The final robustness test considers another particular type of regional spillover and parameter heterogeneity. Budapest could be considered a dominant unit which influences house prices in every other region, but without any feedback from the rest of the country to Budapest. In this model the parameter of Budapest house prices is allowed to vary across regions in both the long-run and the short-run price equations. To save space only impulse responses to a Budapest price shock are reported (see Graph 11). This model confirms significant spillovers to Pest County, which appear with a lag of approximately two years. The impact on prices and quantities in Pest County is around one-fifth of the reaction within Budapest. For the remaining regions, the picture is mixed, some respond quite strongly while others insignificantly.

Overall, the robustness checks regarding neighbourhood effects confirm that regional spillovers exist, and that they are economically significant, although their exact size can be estimated with some uncertainty. In the case of Budapest-specific shocks, spillovers are the strongest in, but not limited to Pest County. There is evidence of heterogeneity in regional elasticities which may warrant further investigation.

Graph 11. **Robustness check: the response of house prices and housing stock to a temporary 10% rise of Budapest house prices, with heterogeneous elasticities to Budapest prices (%)**



Note: horizontal axis denotes time in years; vertical axis is measured in percentages. Mean response (blue solid line) and 95% confidence intervals (red dashed line), based on 1000 bootstrap repetitions of the estimated model. Responses for the rest of Hungary are averages for the remaining six regions; orange dotted lines show the minimum and maximum responses across these regions.

6. CONCLUSION

Regional heterogeneity and spatial spillovers are important but less understood aspects of housing markets. Similarly, the literature often analyses house prices and housing supply separately, downplaying the feedback between quantities and prices. This paper proposed a simple model of regional housing markets, which features spatial spillovers in prices and endogenous housing supply. The model is empirically flexible and it can accommodate various theories of supply. The model was estimated on regional housing data in Hungary.

The model reproduces key elasticities identified in previous aggregate studies. It also allows identifying significant regional spillovers. These spillovers are of relevance in the case of Hungary, where Budapest house prices already appear overvalued. Depending on specification, a 10% overvaluation of Budapest properties leads to a 2-5% rise in the prices of housing in surrounding Pest county, and smaller but statistically significant price increases in other, more distant regions. These higher prices also induce construction activity outside Budapest as well; Budapest and the rest of the country each account for broadly half of the aggregate response of construction over the medium term.

The main policy takeaway from this study is that the housing markets of individual regions should not be viewed in isolation. Spillover effects are significant and local house price misalignments can have visible macroeconomic consequences. Enriching the surveillance of macroeconomic imbalances with the regional dimension can be useful because aggregate house price indexes can blur large underlying heterogeneity. Region-specific house price misalignments might call for spatially differentiated macroprudential regulations (e.g. loan-to-value ratios). Furthermore, the regional analysis calls attention to the potential role of international investors in the housing market, a relatively new and poorly understood development. Finally, regional spillovers induce suburbanisation (mainly around Budapest), which boosts construction and economic activity in neighbouring municipalities, but may also increase traffic congestion, pollution and spatial mismatch between the supply and demand of public services (see Kovács-Tosics, 2014). These challenges can be addressed by coordinated interventions in several policy areas.

In this paper the heterogeneity of spillovers arose only through the weighting matrix, which was based on interregional migration flows. However, some robustness checks hint at heterogeneity in elasticities as well. In addition, asymmetries may also be present on the housing market, i.e. positive and negative shocks can have different effects, for example due to the downward rigidity of the housing stock. The model could be augmented with a separate equation for population, because migration is influenced by differences in regional labour market conditions and house prices. The equation of housing supply could be enriched to take into account local regulatory constraints (e.g. regarding building zones) and construction costs, for which no regional breakdown is currently available. The Hungarian government has applied various policy measures to support homebuyers, including subsidised mortgages and grants depending on the number of children. The regional heterogeneity in the take-up of government support could help in estimating the impact of these measures. Finally, region-specific exogenous shocks (e.g. different exposure to the revaluation of foreign currency denominated household debt) could alleviate remaining concerns of endogeneity in the estimation. These are fruitful areas of further research as more data become available.

REFERENCES

Aron, J., J. V. Duca, J. Muellbauer, K. Murata and A. Murphy (2012). Credit, Housing Collateral, and Consumption: Evidence from Japan, the U.K., and the U.S. Review of Income and Wealth 58: 397-423.

Badarinza, C. and T. Ramadorai (2018). Home Away from Home? Foreign Demand and London House Prices. Journal of Financial Economics 130: 532-555.

Bailey, N., S. Holly and M. H. Pesaran (2016). A Two-Stage Approach to Spatio-Temporal Analysis with Strong and Weak Cross-Sectional Dependence. Journal of Applied Econometrics 31: 249-280.

Baltagi, B., Q. Feng and C. Kao (2012). A Lagrange Multiplier Test for Cross-Sectional Dependence in a Fixed Effects Panel Data Model. Journal of Econometrics 170: 164-177.

Beenstock, M. and D. Felsenstein (2010). Spatial Error Correction and Cointegration in Nonstationary Panel Data: Regional House Prices in Israel. Journal of Geographical Systems 10: 189-206.

Beenstock, M. and D. Felsenstein (2007). Spatial Vector Autoregressions. Spatial Economic Analysis 2: 167-196.

Békés, G., Á., Horváth and Z. Sápi (2016). Lakóingatlan-árak és települési különbségek. MTA Discussion Papers, 2016/21, Hungarian Academy of Sciences, Budapest. (in Hungarian)

Berki, T. and T. Szendrei (2017). The Cyclical Position of Housing Prices – A VECM Approach for Hungary. MNB Occasional Papers 126, Magyar Nemzeti Bank, Budapest.

Brady, R. R. (2014). The Spatial Diffusion of Regional Housing Prices across U.S. States. Regional Science and Urban Economics 46: 150-166.

Bricongne, J. C., A. Turrini and P. Pontuch (2019, forthcoming). Assessing House Prices: Insights from HouseLev, A Dataset of Price Levels Estimates. European Commission Discussion Paper, Brussels.

Chudik, A., M. H. Pesaran and E. Tosetti (2011). Weak and Strong Cross-Section Dependence and Estimation of Large Panels. Econometrics Journal 14: 45-90.

Claeys, G., K. Efstathiou and D. Schoenmaker (2017). Spotting Excessive Regional House Price Growth and What to Do About It. Bruegel Policy Contribution, no. 26.

Corrado, L. and B. Fingleton (2012). Where Is the Economics in Spatial Econometrics? Journal of Regional Science 52: 210-239.

Farkas, M. (2011). Housing Demand and Demographic Trends: Evidence from Hungary. M.A. Thesis, Central European University, Department of Economics, Budapest.

Gattini, L. and P. Hiebert (2010). Forecasting and Assessing Euro Area House Prices through the Lens of Key Fundamentals. ECB Working Paper, No. 1249, European Central Bank, Frankfurt.

Gonzalo, J. (2004). Five Alternative Methods of Estimating Long-Run Equilibrium Relationships. Journal of Econometrics 60: 203-233.

Grimes, A. and A. Aitken (2010). Housing Supply, Land Costs and Price Adjustment. Real Estate Economics 38: 325-353.

Haurin, D. R., P. H. Hendershott and D. Kim (2003). The Impact of Real Rents and Wages on Household Formation. Review of Economics and Statistics 75: 284-293.

Holly, S., M. H. Pesharan and T. Yamagata (2011). The Spatial and Temporal Diffusion of House Prices in the UK. Journal of Urban Economics 69: 2-23.

Hviid, S. J. (2017). A Regional Model of the Danish Housing Market, Danmarks Nationalbank Working Paper, No. 121, Copenhagen.

Horváth, Á. (2008): Increasing House Prices since 1995: Causes and Measurement. Ph.D. Thesis, Corvinus University of Budapest, Economics Doctoral School, Budapest.

Illés, S. and G. Michalkó (2008). From Tourism to Migration: Spatial Inequality of Foreign Property Owners in Hungary. In: Á. Kertész and Z. Kovács (eds), Dimensions and Trends in Hungarian Geography, Dedicated to the 31st International Geographical Congress, Tunis, 12-15 August 2008. Hungarian Academy of Sciences Geographical Research Institute, Budapest: 207-222.

Jancsik, A., G. Michalkó and M. Csernyik (2018): Megosztás megosztottság nélkül – az Airbnb és a budapesti szálláshelypiac átalakulása. Közgazdasági Szemle 65: 259-286. (in Hungarian)

Kao, C. and M. H. Chiang (2000). On the Estimation and Inference of a Cointegrated Regression in Panel Data. Nonstationary Panels, Panel Cointegration, and Dynamic Panels 15: 179-222.

Kiss G. and G. Vadas (2007). The Role of the Housing Market in Monetary Transmission – Evidence from Hungary. European Journal of Housing Policy 7: 299-317.

Liao, W. C., D. Zhao, L. P. Lim and G. K. M. Wong (2015). Foreign Liquidity to Real Estate Market: Ripple Effect and Housing Price Dynamics. Urban Studies 52: 138-158.

Lee, K. O. and G. Painter (2013). What Happens to Household Formation in a Recession? Journal of Urban Economics 76: 93-109.

Mankiw, N. G. and D. N. Weil (1989). The Baby Boom, the Baby Bust, and the Housing Market. Regional Science and Urban Economics 19: 235-258.

Mayer, C. J. and C. T. Somerville (2000). Residential Construction: Using the Urban Growth Model to Estimate Housing Supply. Journal of Urban Economics 48: 85-109.

Meen, G. (2016). Spatial Housing Economics: A Survey. Urban Studies 53: 1987-2003.

Meen, G. (2000). Housing Cycles and Efficiency. Scottish Journal of Political Economy 47: 114-140.

MNB (2018). Housing Market Report – May 2018. Magyar Nemzeti Bank, Budapest.

Muellbauer, J. (2012). When is a Housing Market Overheated Enough to Threaten Stability? In: Heath, A., F. Packer and C. Windsor (eds), Property Markets and Financial Stability, Conference Proceedings. Reserve Bank of Australia, Sydney: 73-105.

Muellbauer, J. and A. Murphy (2008). Housing Markets and the Economy: The Assessment. Oxford Review of Economic Policy 24: 1-33.

Oikarinen, E. (2004). The Diffusion of Housing Price Movements from Center to Surrounding Areas. Journal of Housing Research 15: 3-28.

Pesaran, M. H. (2015). Testing Weak Cross-Sectional Dependence in Large Panels. Econometric Reviews 34: 1089-1117.

Pesaran, M. H. (2007). A Simple Panel Unit Root Test in the Presence of Cross-Section Dependence. Journal of Applied Econometrics 22: 265-312.

Pesaran, M. H. and Y. Shin (1998). An Autoregressive Distributed-Lag Modelling Approach to Cointegration Analysis. In: S. Strøm (ed.), Econometrics and Economic Theory in the 20th Century. Cambridge University Press: 371-413.

Philiponnet, N. and A. Turrini (2017). Assessing House Price Developments in the EU. European Economy Discussion Paper 048, European Commission, Brussels.

Poterba, J. M. (1984). Tax Subsidies to Owner-Occupied Housing: An Asset-Market Approach. Quarterly Journal of Economics 99: 729-752.

Székely, J. (2008). Reasons of Local Housing Market Price Differences in Hungary. Hungarian Statistical Review 86: 174-188.

Takáts, E. (2012). Aging and House Prices. Journal of Housing Economics 21: 131-141.

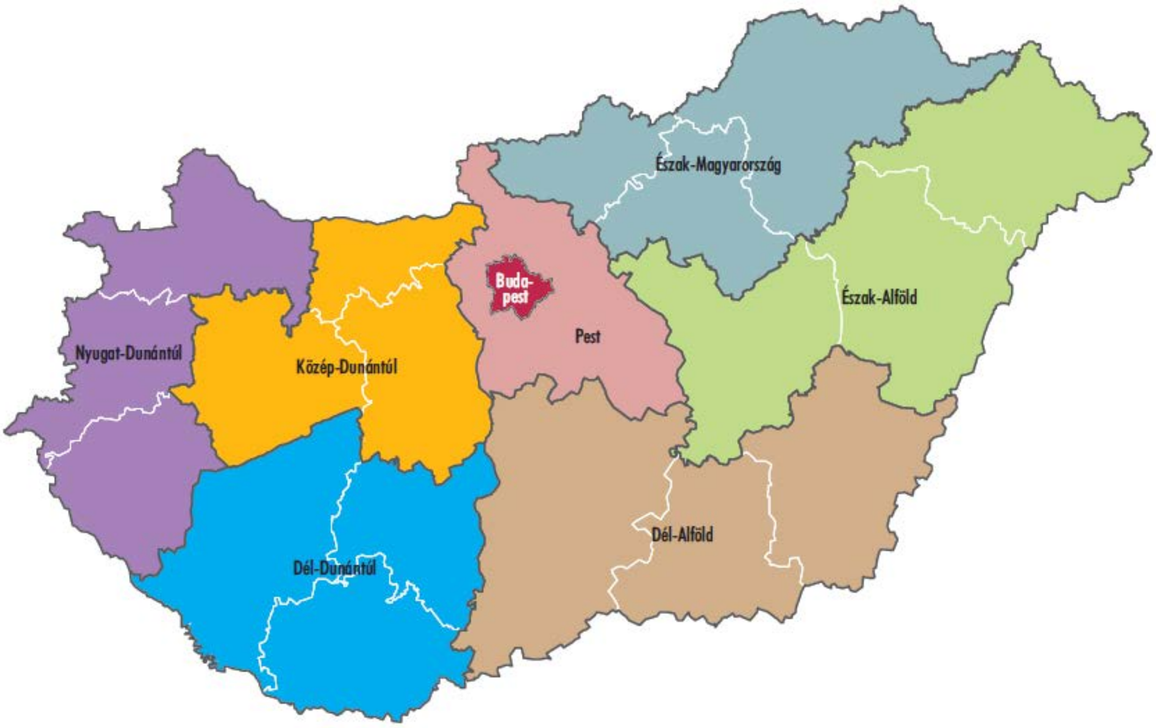
Teye, A. L., M. Knoppel, J. de Haan and M. G. Elsinga (2017). Amsterdam House Price Ripple Effects in the Netherlands. Journal of European Real Estate Research 10: 331-345.

Vadas, G. (2009). The Housing Subsidy Scheme and Households' Wealth in Hungary: Urban Legends and Facts. European Journal of Housing Policy 9: 1-24.

Verner, E. and Gy. Gyöngyösi (2018): Household Debt Revaluation and the Real Economy: Evidence from a Foreign Currency Debt Crisis. Mimeo, 8 October 2018. Available at SSRN: https://dx.doi.org/10.2139/ssrn.3241309

ANNEX

Graph 12. **Statistical regions of Hungary**



Source: Hungarian Central Statistical Office

EUROPEAN ECONOMY DISCUSSION PAPERS

European Economy Discussion Papers can be accessed and downloaded free of charge from the following address:
https://ec.europa.eu/info/publications/economic-and-financial-affairs-publications_en?field_eurovoc_taxonomy_target_id_selective=All&field_core_nal_countries_tid_selective=All&field_core_date_published_value[value][year]=All&field_core_tags_tid_i18n=22617.

Titles published before July 2015 under the Economic Papers series can be accessed and downloaded free of charge from:
http://ec.europa.eu/economy_finance/publications/economic_paper/index_en.htm.

GETTING IN TOUCH WITH THE EU

In person

All over the European Union there are hundreds of Europe Direct Information Centres. You can find the address of the centre nearest you at: http://europa.eu/contact.

On the phone or by e-mail

Europe Direct is a service that answers your questions about the European Union. You can contact this service:

- by freephone: 00 800 6 7 8 9 10 11 (certain operators may charge for these calls),
- at the following standard number: +32 22999696 or
- by electronic mail via: http://europa.eu/contact.

FINDING INFORMATION ABOUT THE EU

Online

Information about the European Union in all the official languages of the EU is available on the Europa website at: http://europa.eu.

EU Publications

You can download or order free and priced EU publications from EU Bookshop at: http://publications.europa.eu/bookshop. Multiple copies of free publications may be obtained by contacting Europe Direct or your local information centre (see http://europa.eu/contact).

EU law and related documents

For access to legal information from the EU, including all EU law since 1951 in all the official language versions, go to EUR-Lex at: http://eur-lex.europa.eu.

Open data from the EU

The EU Open Data Portal (http://data.europa.eu/euodp/en/data) provides access to datasets from the EU. Data can be downloaded and reused for free, both for commercial and non-commercial purposes.

Publications Office
of the European Union