

Financial Infrastructure and House Prices

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Abstract

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Abstract We argue that banks operating in a local market possess better information about the local housing market than do non-local banks. Possessing this information may influence their willingness to grant loans to house buyers and the specifics of the loan terms, which in turn may affect house prices because credit facilitation makes the housing market more efficient. Using a panel data set covering a period from 1993 to 2007 and involving 274 municipalities in Sweden, we establish a positive causal influence of local bank presence on local house prices. There are significant spatial and spillover effects – that is, banks in a municipality affect the housing markets in neighboring municipalities, although to a lesser extent than in their own municipality. Similar results are obtained through a gravity model. The results are robust over time and municipality size.

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Introduction

An old adage in the real estate business states that the value of a house depends on the following three factors: “location, location, location.” Although this may not be the whole story, it is certainly true that the immobility of real estate means that its location has a large impact on its value. Thus, it is important for all agents in the housing business - buyers, sellers, brokers, developers, and financial intermediaries (the focus of this paper) - to possess a thorough understanding of the local market.

We propose - and then empirically test - a hypothesis stating that the presence and extent of local financial infrastructure may influence house values. The underlying notion is that local financial intermediaries, e.g. banks, are better informed than non-local ones about the local housing market and, thus, better equipped to make sound financial decisions, for example, regarding whether to grant loans and under what terms. We argue that these financial decisions may, in themselves, affect house values.

The aim of this paper is to investigate whether such a causal relationship between the proximity to financial infrastructure and house prices exists. We test this by employing an error correction model with instrument variables. The results suggest that local financial infrastructure indeed influences the house prices. Spatial and spillover effects - the influences from one local region on nearby regions – turn out to be important. We also find support for a “crystal ball” effect in that local banks seem to foresee future development of their local housing market more accurately than non-local ones.

The remainder of this paper is organized into five sections. The underlying theoretical foundations, drawn from two different literature fields, are presented in the following section. In the following section, we present the econometric modeling approach. The third section describes the data. The empirical analyses are presented in the fourth section. Results and conclusions are summarized in the fifth and final section.

Literature Review and Hypotheses

The present paper draws from two branches in the literature. The first comprises studies that investigate the relationship between bank lending and house prices typically aiming to understand the causal direction. Several theoretical reasons for the existence of a positive and bidirectional relationship between bank lending and house prices have been put forth (Oikarinen 2009). On the one hand, increased bank lending may lead to higher house values as more credit in the market may lower interest rates, thus making future housing benefits more valuable today. Furthermore, it may reduce credit constraints, and/or it may serve as a signal that borrowers have positive beliefs about variables that influence the housing market (e.g. future income levels). On the other hand, increased house values may influence bank lending as they increase the wealth of house owners, which may reduce their borrowing constraints. They may also influence households' lifetime wealth, which may affect current spending and borrowing, and they may increase the "wealth" of banks, which may enhance the willingness and feasibility of banks to grant new loans.

Understanding the relationship between lending and house prices is important, not least, from a policy perspective. For instance, it may help policymakers in determining the feasibility of cooling an overheated housing market through policies that restrict credit supply. Even though several empirical studies have been conducted on this topic, there remains a lack of consensus. For instance, Gerlach and Peng (2005) found that house prices affect lending, but they did not find that bank lending affects house prices. Brissimis and Vlassopoulos (2009) found similar results in the long term but found that the causal effect is bidirectional in the short term. This bidirectional effect has also been identified in other studies (e.g. Liang and Cao 2007; Goodhart and Hofmann 2008; Koetter and Poghosyan 2010; Gimeno and Martínez-Carrascal 2010). The lack of consensus may partly be due to that different studies rely on data from different nations, leading to large variations in institutional settings etc.

For instance, Oikarinen (2009) found that a bidirectional effect is more pronounced in deregulated markets.

The second literature branch of importance here focuses on the financial infrastructure's ability to enhance economic growth. This positive relationship has been well established (e.g. Goldsmith 1969). However, the direction of causality is not trivial to determine (Levine and Zervos 1998; Levine 2005). Economic growth in a region may make it more attractive for banks and other financial actors to locate there and, hence, it seems likely that local economic growth may have a positive impact on local financial infrastructure.

There are theoretical reasons for anticipating the existence of causality also in the opposite direction. For the purposes of this paper, we are particularly interested in two mechanisms that may explain this causality. First, the financial sector may improve the accuracy of information available about firms, managers, and economic conditions, see, e.g. Greenwood and Jovanovic (1990). A recent and growing literature focuses on that distance hinders the collection of, so called, soft information – thereby establishing a link between (inverse) distance and information (Boot 2000; Petersen and Rajan 2002; Berger and Udell 2002; Degryse and Ongena 2005; Alessandrini et.al. 2009; Butler 2008; Agrawal and Hauswald 2010). Second (and highly related), banks may provide monitoring and, thereby, influence borrowers to behave in an economically sound manner (Diamond 1984).

These theoretical mechanisms typically address the impacts on small firms. In the present paper, we apply similar ideas to the case of regional housing markets. This implies that the local presence of banks or other types of financial infrastructure may provide information that would not be available otherwise; that is, a local bank has superior information about the local market and is thereby better equipped to make decisions regarding housing loans. This is thus a parallel to finding positive net present value projects in the firm financing case.

The presence of local financial infrastructure could potentially influence values both upward and downward. The idea would be that there is a true condition of the housing market, about which a local bank is better (but not fully) informed than a non-local bank. If the local bank has a more positive belief about the market than the non-local one, it is more likely to grant a loan, which in turn may further increase house values. However, the opposite should also apply. If the local bank has a more pessimistic market view, it should be less willing to grant loans, which in turn may decrease house values.

Even so, this reasoning may lead to a likely positive causal relationship between local bank presence and house values. For instance, if a local bank does not grant a loan, the borrower may turn to a non-local bank. If the non-local bank grants a loan in such a situation, there is no longer any negative impact on house values from a local bank presence in bad markets. Thus, local banks will increase house values on average¹.

Once a loan is granted, a local bank may provide better monitoring than a non-local bank and, in so doing, have a positive impact on local house values. It is in their best interest that house owners act appropriately so as to increase the value of the collateral.

Taken together, these ideas lead to a series of testable hypotheses. In particular, increasing the presence of local financial infrastructure, which is measured as the number of bank offices or bank employees in a given region, should, on average, increase house values. Similarly, a well-developed financial infrastructure in adjacent regions should also influence house values, but to a lesser extent, i.e., there should be positive spillover effects. The positive impact on house values should be more pronounced when and where the market is booming. In some situations, the increased

¹ Eventually, the non-local bank should understand the adverse selection problem it faces; only customers in bad markets turn to non-local banks. As a response, the non-local bank should be more restrictive in granting new loans. This further strengthens the effect that local bank presence will increase house values.

presence of local infrastructure may reduce house values. If this is the case, this should occur in bad markets. Even so, the average impact should be positive.

The Econometric Model

The objective is to estimate the causal relationship between regional house prices and the presence of bank branches or bank employments in the region. The general model follows the ideas put forth by DiPasquale and Wheaton (1994), Jud and Winkler (2002), and Wilhelmsson (2008). In equilibrium, the reduced form housing price equation in a given period and municipality reflects the demand and the supply of housing services during that period and in that municipality. There are a number of econometric issues that need to be handled. We may have problems with non-stationary time series and simultaneous causality (endogeneity) as well with spatial and temporal autocorrelation.

The equation of main interest is a first-difference model with an error correction term, that is, the short-run equation in an Engel-Granger two-step procedure of an error correction model (EMC). The stochastic version of the model is:

$$\Delta HP_{i,t} = \hat{\alpha} + \hat{\beta}_1 \Delta X_{i,t} + \hat{\beta}_2 \overline{\Delta Bank}_{i,t} + \hat{\beta}_3 \overline{W \Delta HP}_{i,t} + \hat{\beta}_4 \Delta Spillover_{i,t} + \beta_5 T_t + \lambda ECT_{i,t-1} + \varepsilon_{i,t} \quad (1)$$

where HP denotes house prices (natural logarithm), X is a vector of control variables, i.e., income per capita and housing stock per capita, $Bank$ is the number of bank branches or bank employees per capita, and W is a spatial weight matrix. $Spillover$ is designed to capture the impact from financial infrastructure in nearby regions. We employ two different specifications to achieve this as discussed below. ECT is the error correction term, T is a number of binary variables indicating years (capturing trend and omitted variables constant across municipalities such as mortgage interest rate). The subscript i refers to municipality and t to year.

In order to control for non-stationarity in the series, we test for unit root in the individual time series. Earlier literature suggests that at least some of the series should be non-stationary, Roll (2002). We also test for if the non-stationary series are co-integrated. If they are, an error correction model is possible to estimate.

The model is estimated using FGLS with an AR(1) autocorrelation within panels and heteroskedasticity across panels. The GLS method reduces problems of unequal variance, and/or temporal correlation between observations. In the present case, the covariance is unknown, and the parameters need to be estimated, and, thus, we use the FGLS method (see Greene 2008). The major reason for not estimating a fixed effect model is the low within-panel correlation concerning our main variable of interest (*Banks*).

The error correction term (*ECT*) is the residual from the long-run equation where the relationship is estimated in levels (non-stationary data). It is estimated as:

$$ECT_{i,t} = \alpha_0 + \alpha_1 X_{i,t} + \alpha_2 Bank_{i,t} + \alpha_3 WHP_{i,t} + \alpha_4 Spillover_{i,t} + \alpha_5 T_t \quad (2)$$

If *ECT* is stationary, the dependent and the independent variables are co-integrated. We are using two different tests in order to detect for unit root in the variables and in the residual, namely Levin-Lin-Chu (LLC) test (Levin et al, 2002) and Im-Pesaran-Shin (IPS) test (Im et al, 2003). The LLC assumes that all panels have the same autoregressive parameter. IPS relaxes the assumption and allows each panel to have its own autoregressive parameter.

A major concern is the problem of simultaneous causality (endogeneity), that is, we do not know if the bank chooses to locate in regions that they believe will exhibit future positive growth in house prices or if the growth in house prices are a consequence of the presence of banks. Hence, the causality may go in both directions. If so, the error and the independent variables are correlated and the OLS estimates

will be biased. To handle this, we employ an instrument variable two-stage technique where the instruments are assumed to be uncorrelated with the error term but a good proxy for the endogenous bank variable. The second stage in the two-stage procedure is the estimation of equation 1 above. The first stage regression is:

$$\overline{\Delta Bank_{i,t}} = \eta_0 + \eta_1 \Delta Rob_{i,t} + \eta_2 Rob_{i,t} + \eta_3 X_{i,t} \quad (3)$$

where $Rob_{i,t}$ is the number of robberies in bank, post offices, and value transports in municipality i at time t . We use this as an instrument variable arguing that there is little reason to believe that the number of bank robberies has a direct impact on regional house prices, but it is presumably positively correlated with the number of banks. We employ Sargan's test to test for validity in the instruments and instrument relevance has been tested by the F-statistics in the first stage regression, see Murray (2006) and Basile (2008) for a discussion how to avoid invalid instruments and coping with weak instruments.

Finally, we have potential problems with spatial dependency and spatial spillover. In order to handle spatial dependency, a spatial lag model is estimated. Spillover from banks in one region to another region is tested using two different spillover representations; a spillover based on the number of bank branches/employees in the rest of the county, and a gravity representation measuring bank branches/employees in all of Sweden weighted by the (inverse) distance between municipalities. The spatial lag model cannot be estimated with the FGLS method because the spatially lagged dependent variable is endogenous. To reduce this problem, we adopt the IV approach put forth by Kelejian and Prucha (1998). This method has been empirically used in previous research (e.g. Buettner 2001; Solé Ollé 2003; Usai and Paci 2003) and in more recent research (Basile 2008). We use neighboring municipalities' average

income per capita, housing stock and population size as instruments in the model. The expected value of the spatially lagged house price difference is given by:

$$\overline{W\Delta HP}_{i,t} = \gamma_0 + \gamma_1 WX_{i,t} \quad (4)$$

The spatial weight matrix, W , is defined by the inverse distance between municipalities and is row standardized.

In addition, we construct variables to test the hypothesis regarding spillover effects across municipality borders (*Spillover* in equation 1). The first two constructed variables measure the average number of bank offices (VBO) and bank employees (VBE) in the rest of the county thereby capturing the size of the financial infrastructure in adjacent municipalities. These variables are defined by the following equations:

$$VBO_i = (\sum_{j \in k} BO_j - BO_i) / (m_k - 1) \text{ and } VBE_i = (\sum_{j \in k} BE_j - BE_i) / (m_k - 1) \quad (5)$$

where BO_j and BE_j , respectively, denote the number of bank offices and bank employees in municipality j , and m_k is the number of municipalities in county k in which municipality i is located.

As alternative measures, we construct two additional variables, denoted GBO and GBE , by using information about the geographical location of the municipality. These variables summarize the distance of each municipality to all bank offices (GBO) and bank employees (GBE) in other municipalities, using a gravity representation. The variables are defined by the following equation:

$$GBO_i = \sum_{j \neq i} (BO_j / d_{ij}) \text{ and } GBE_i = \sum_{j \neq i} (BE_j / d_{ij}) \quad (6)$$

where d_{ij} is the distance between communities i and j .

Data and Descriptive Statistics

We employ a panel data set consisting of 274 (of 284²) municipalities in Sweden over a period of 15 years (1993–2007). Thus, the investigated period began at the very end of the real estate crisis that started in 1990 in Sweden. In the middle of the 1980s, Sweden experienced huge credit expansion. Rents and real estate prices increased by more than 240% and 790%, respectively, between 1980 and 1989. Construction increased, usually with 100% debt, and vacancy rates were low. One consequence of the 1990/1991 Swedish tax reform was a reduction in the deductibility of interest rates from 50% to 30%. A rapid drop in demand for real estate around 1990 put Sweden and its banks under severe stress, and Sweden experienced a real interest rate shock. In the fall of 1992, policymakers decided to implement two acute measures; a floating currency and an inflation target for the central bank. In 1995, Sweden became a member of the European Union (EU). Although the real estate crises ended in 1994, their effects were felt in Sweden's real estate market and banking industry for a number of years. In 2000, a new fiscal policy framework was introduced at the same time that the central bank became independent. The central bank also began producing financial stability reports at this time. Moreover, Basel II³ was introduced, which affects bank lending to households.

The Swedish banking market is dominated by four large banks, which account for approximately three quarters of the market. The remaining market is served partly by Swedish-owned banks and partly by foreign banks. The former include banks with different ownership structures and market strategies. Notably, some saving banks have been transformed into banking companies, and banks have been established that serve a niche market. These niche banks often distribute their services via the Internet or by telephone and, thus, have a minimum number of bank offices. This group has

² During the study period, some new municipalities have merged. These municipalities are excluded from the data.

³ Basel II is a set of recommendations for bank laws meant to establish an international standard.

increased rapidly in number since the early 1990s. Foreign banks have been allowed to set up subsidiaries in Sweden since 1986 and to open branch offices since 1990. By 2010, there were 29 foreign banks operating in Sweden. Although most of them do not target the home mortgage market, there are exceptions.

Almost 70% of Swedish households are homeowners. Of these, 81% have a mortgage, with their home being used as collateral. Typically, mortgages are issued by mortgage institutes that often are closely tied to the banks, Swedish Bankers' Association (2011).

The basic information in the data set regards house prices and the numbers of bank offices and bank employees in each municipality. We also have information concerning the size of the municipality, measured as the number of inhabitants, the housing stock of single-family houses, and the income per capita.

Table 1. Definition and descriptive statistics

Variable	Definition	Average	Standard deviation
House price	Nominal average house prices	788.41	631.20
Income	Disposable income per capita	167.62	32.87
Bank branches	No. of establishments in the sector financial services	15.55	51.41
Bank employment	No. of employees in the sector financial services	216.92	1523.70
Housing stock	Housing stock per capita	0.2944	0.0748
Population	No. of inhabitants	31738.74	58728.70
VBO	No. of establishments in the financial sector in the rest of county	15.34	11.29
VBE	No. of employees in the financial sector in the rest of county	213.69	323.01
GBO	No. of establishments in the financial sector in the rest of country weighted by distance	0.0266	0.0358
GBE	No. of employees in the financial sector in the rest of country weighted by distance	0.4600	1.0359

The descriptive statistics in Table 1 show that the average house price in Sweden over the 15-year period is 788,000 SEK, but the standard deviation is relatively high.⁴ The standard deviation compared with the mean for income per capita is much lower. The average income per capita equals 160,000 SEK, with a standard deviation of 33,000. Substantial variation exists in the numbers of bank offices and bank employees. The average number of bank offices among the municipalities in Sweden is only 15 (with 217 employees), but the standard deviation is as high as 51 (with 1524 employees). The average number of inhabitants is approximately 31,000, with a coefficient of variation of about 2.

As mentioned, we use the LLC-test and IPS-test to test the estimates for unit root. For all variables, the unit root test is performed on the natural logarithm. The results are given in Table 2. The results indicate that the variables house price and income per

⁴ Here we use average house price. We have also tested using quality adjusted house prices, but there is no change in results. As the time series is much shorter for the quality adjusted house prices we are utilizing the non-quality adjusted average house price in the subsequent empirical analysis.

capita are not stationary. Both variables measuring financial infrastructure (bank branches and bank employments) are integrated of order zero, that is, stationary.

Table 2. Unit root test (t-values).

	Levin-Lin-Chu	Im-Pesaran-Shin	Integrated of order
House price	27.7962	40.1933	1
Δ House price	-39.2681	-33.3120	
Income per capita	-16.4118	7.9963	1
Δ Income per capita	-27.3627	-12.6086	
Housing stock per capita	-16.5342	-2.5793	0
Δ Housing stock per capita	-28.3784	-19.0506	
Bank branches	-19.3161	-14.5130	0
Δ Bank branches	-39.7685	-30.1855	
Bank employment	-10.8940	-2.6539	0
Δ Bank employment	-35.2270	-29.4015	

Note. All variables are in the form of natural logarithm.

Over the studied time period, the nominal house prices increased by a remarkable 150% on average. However, the average number of local bank offices was almost the same in 2007 as it was in 1993. Based on these data, it does not seem as the number of bank offices drives house prices.

Econometric Analysis

The main objective is to estimate the first-difference model with error correction given by equation (1). In order to do this, we first need to estimate equations (2)-(6). Table 3 presents the results from the FGLS models estimating the long-run equation (2).

Model A1 and C1 use bank establishments per capita as independent variable while B1 and D1 use bank employments per capita as independent variable. Model A1 and B1 use average financial infrastructure in the county as spillover representation. C1 and D1 use gravity spillover representation.

In these long-run models, house prices are a function of demand and supply. We employ income per capita as the only demand variable and housing stock per capita as a measurement of supply. All models are random effects models. We use an instrument variable (IV) approach so as to make the bank per capita variables

exogenous, see equation (3). To test for overidentification we employ the Sargan's test, which tests the validity of instrumental variables. The null hypothesis is that the instruments are valid; that is, we do not want to reject the null hypothesis.

Table 3. Error-Correction Model: Long-run equation.

	Model A1	Model B1	Model C1	Model D1
Income per capita	2.1453 (94.22)	2.0993 (88.62)	2.2396 (89.95)	2.2847 (88.77)
Stock per capita	-0.8020 (-66.36)	-0.6805 (-53.44)	-0.8187 (-72.05)	-0.7782 (-63.83)
Bank Establishments	0.0904 (23.64)	-	0.0371 (8.52)	-
Employments	-	0.0894 (27.82)		0.0330 (8.41)
Spillover (Bank) GBO/GBE	-	-	-0.0065 (-1.09)	-0.0181 (-4.04)
VBO/VBE	0.0585 (10.72)	0.0383 (11.58)	-	-
Spatial lag (House prices)	0.1420 (50.47)	0.1443 (53.01)	0.1561 (54.34)	0.1621 (60.75)
Constant	-5.2397 (-36.25)	-5.2888 (-39.93)	-6.7415 (-33.02)	-7.1588 (-39.52)
AIC	-2139.64	-2123.35	-2096.86	-2101.36
BIC	-343.62	-320.15	-293.67	-298.16
Co-integration				
Im-Pesaran-Shin (p-value)	0.0000	0.0000	0.0000	0.0000
Levin-Lin-Chu (p-value)	0.0000	0.0015	0.0026	0.0072

Note. Dependent variable: natural logarithm of house prices. We are controlling for spatial and temporal correlation as well as heteroskedasticity by estimating a feasible GLS. Time dummies are not shown in the table. t-values within brackets. The spatial matrix and the spatial lag variable have been estimated using Stata-code spwmatrix and splagvar (P. Wilner Jeanty's. 2010). We are using two tests (Im-Pesaran-Shin och Levin-Lin-Chu) for unit root in the error term in order to test for co-integration. The result shows that the variables are co-integrated. AIC and BIC are used to compare the model specifications. Model A1 is preferred.

The estimates from the IV models show that income per capita has a positive impact on house prices, as expected. Moreover, the size of the housing stock is inversely related to house prices. The estimates are robust across the different specifications. The numbers of establishments and employees per capita within the financial services industry are positively related to house prices.

The estimates concerning the spatial lags are highly significant different from zero indication that there is a spillover across municipalities. When measured as averages in the county (model A1 and B1) the spillover effects from bank presence in other municipalities are positive and significant. When measured using a gravity approach, the spillovers are not significant (model C1 and D1). The main reason for this is that the spatial lags pick up most of the spatial dependency that may exist in the data. If we run the model without the spatial lags, the spillover estimates are positive and statistically significant different from zero (not shown).

We are using AIC and BIC in order to choose between the different specifications. This suggests that specification A1 is preferred. The unit root test confirms that house prices and income per capita are co-integrated.

As noted, our primary objective is to estimate equation (1). The results from that model are presented in Table 4. Four different specifications are presented (Model A2-D2).

Table 4. Error-Correction Model: Short-run equation.

	Model A2	Model B2	Model C2	Model D2
Δ Income per capita	0.9193 (5.35)	0.5062 (3.22)	1.0568 (6.15)	0.5452 (3.49)
Δ Stock per capita	-0.9892 (-4.23)	-0.7828 (-4.83)	-0.8481 (-3.62)	-0.4885 (-2.80)
Bank				
Δ Establishments	0.4139 (2.23)	-	0.3328 (1.82)	-
Δ Employments	-	0.9093 (7.85)	-	0.8270 (7.08)
Spillover (Bank)				
Δ GBO/ Δ GBE	-	-	-0.2526	-0.3948

			(-4.37)	(-4.54)
$\Delta VBO/\Delta VBE$	0.0008	0.0073	-	-
	(0.04)	(0.36)		
Δ Spatial lag (House prices)	1.0296	0.8654	1.1464	0.9107
	(11.85)	(10.11)	(13.12)	(10.88)
ECT(t-1)	-0.0236	-0.0187	-0.0263	-0.0243
	(-4.82)	(-3.87)	(-5.49)	(-5.12)
Constant	-0.037	0.0170	-0.0357	0.0177
	(-3.20)	(1.44)	(-3.40)	(1.52)
AIC	-7651.59	-7730.98	-7701.35	-7750.36
BIC	-5881.15	-5953.45	-5923.82	-5972.83
Sargan test	0.9263	0.6049	0.9557	0.5937
Serial correlation	0.000	0.000	0.000	0.000
Heteroscedasticity	0.000	0.000	0.000	0.000
Hausman test	0.000	0.013	0.014	0.000

Note. Dependent variable: the change of natural logarithm of house prices. We are controlling for spatial and temporal correlation as well as heteroskedasticity by estimating a feasible GLS. Time dummies are not shown in the table. t-values within brackets. The spatial matrix and the spatial lag variable have been estimated using Stata-code spwmatrix and splayvar (P.Wilner Jeanty's. 2010). Hausman specification test is used to test whether fixed effect or random effect model is preferred. A random effect model is preferred. The test for heteroskedasticity test is a likelihood-ratio test where the null hypothesis is homoskedasticity. Wooldridge test for serial correlation is used where the null hypothesis is no autocorrelation. Sargan's test is a test of the exogeneity of one or more instruments. Under the null hypothesis the instruments are valid. AIC and BIC are used to compare the model specifications. Model D2 is preferred.

Most of the estimates have the expected sign and are highly significant. The spatial lag variable is positive and significant in all specifications indicating that spatial dependency is present. The spillover effects turn out as negative or not significant. It seems that the spatial lag variable captures the spatial dependencies better than the spatial bank variables; VBO, VBE, GBO and GBE. As expected, the estimates concerning the error correction term are negative and significant in all different specifications. However, it seems that the adjustment speed is slow. The Sargan test shows that we cannot reject the null hypothesis, thus suggesting that the instruments are valid.

The main result is that the bank variable is positive and significant in all model specifications (not significant at the 5% level in Model C2). This indicates a positive impact from local bank presence on house prices. Based on AIC and BIC, model D2 is preferred. That model suggests that, if the change in the number of bank

establishments per capita increases by 1%, the change in house prices are expected to increase by 0.4%.

Having established a causal effect on house prices from local bank presence we proceed by analyzing whether the estimates vary over time and/or across space (Tables 5 and 6). First, we test if the estimated effects were higher or lower after the year 2000. The reason to suspect a difference before and after the year 2000 is that Sweden then introduced a number of new policies and institutional changes, as previously described. However, our approach may not distinguish between effects due to, say, the general business cycle and effects due to new fiscal policies. Second, because the effects may vary owing to municipality size, we test if the effects are higher or lower in municipalities with a population of fewer than 10,000 inhabitants.

Table 5. Error-Correction Model: Short-run equation. Before and after 2000.

	Model A3	Model B3	Model C3	Model D3
Δ Income per capita	0.9303 (5.43)	0.5061 (3.23)	0.9666 (5.66)	0.5167 (3.31)
Δ Stock per capita	-0.9775 (-4.18)	-0.7208 (-4.38)	-1.0224 (-4.40)	-0.7330 (-4.46)
Bank				
Δ Establishments	0.5676 (2.56)	-	0.5310 (2.41)	-
Δ Establishments-2000	-0.2349 (-1.27)	-	-0.2448 (-1.33)	-
Δ Employments	-	1.1246 (6.22)	-	1.1079 (6.14)
Δ Employments-2000	-	-0.2809 (-1.58)	-	-0.2651 (-1.50)
Δ Spatial lag (House prices)	1.0275 (12.07)	0.8669 (10.32)	1.0385 (12.23)	0.8687 (10.37)
ECT(t-1)	-0.0227 (-4.59)	-0.0168 (-3.42)	-0.0225 (-4.69)	-0.0196 (-4.08)

Constant	-0.0338 (-3.23)	0.0154 (1.31)	-0.0353 (-3.38)	0.0149 (1.27)
AIC	-7652.89	-7733.01	-7686.92	-7736.82
BIC	-5882.45	-5955.48	-5909.39	-5959.30

Note. Dependent variable: the change of natural logarithm of house prices. We are controlling for spatial and temporal correlation as well as heteroskedasticity by estimating a feasible GLS. Time dummies are not shown in the table. t-values within brackets. The spatial matrix and the spatial lag variable have been estimated using Stata-code spwmatrix and splayvar (P.Wilner Jeanty's. 2010). AIC and BIC are used to compare the model specifications. Model D3 is preferred.

Table 5 shows the results from the regressions examining parameter heterogeneity before and after the year 2000. According to the AIC and BIC, model D3 is preferred.

The estimates concerning income and housing stock are robust. The number of bank employees in the financial services industry has a positive and significant effect on house prices over this period. The effect is somewhat (around 0.25%) lower between 2000 and 2007 than between 1993 and 1999. However, the difference between time periods is not significant. Thus, the results suggest that the effect is quite constant over time.

We test whether there is parameter heterogeneity across regions by examining interactions between the bank variables and a dummy variable indicating if the region has fewer than 10,000 inhabitants. The results are shown in Table 6.

Table 6. Error-Correction Model: Short-run equation. Small and large region.

	Model A4	Model B4	Model C4	Model D4
Δ Income per capita	0.9144 (5.32)	0.5143 (3.27)	0.9509 (5.55)	0.5248 (3.34)
Δ Stock per capita	-0.9848 (-4.21)	-0.7802 (-4.83)	-1.0312 (-4.43)	-0.7895 (-4.89)
Bank				
Δ Establishments	0.4165 (2.25)	-	0.3729 (2.03)	-
Δ Establishments-small	0.0206 (0.52)	-	0.0198 (0.50)	-
Δ Employments	-	0.8930 (7.52)	-	0.8886 (7.48)
Δ Employments-small	-	0.0260 (0.39)	-	0.0268 (0.40)
Δ Spatial lag (House prices)	1.0291 (12.03)	0.8713 (10.32)	1.0406 (12.20)	0.8731 (10.37)
ECT(t-1)	-0.0236 (-4.82)	-0.0187 (-3.88)	-0.0235 (-4.93)	-0.0213 (-4.50)
Constant	-0.0334	0.0160	-0.0349	0.0154

	(-3.18)	(1.36)	(-3.34)	(1.31)
AIC	-7651.81	-7731.01	-7685.72	-7735.06
BIC	-5881.37	-5953.49	-5908.19	-5957.53

Note. Dependent variable: the change of natural logarithm of house prices. We are controlling for spatial and temporal correlation as well as heteroskedasticity by estimating a feasible GLS. Time dummies are not shown in the table. t-values within brackets. The spatial matrix and the spatial lag variable have been estimated using Stata-code spwmatrix and splayvar (P.Wilner Jeanty's. 2010). AIC and BIC are used to compare the model specifications. Model D4 is preferred.

All models in Table 6 yield similar qualitative results; there is a positive and highly significant impact of banks on house prices, but the effect does not differ between small and large municipalities.

Finally, as shown in Table 7, we test whether the impact of financial infrastructure on housing prices is asymmetric in the sense that it differs for housing markets in a boom or bust state. Our hypothesis is that the presence of banks should have a positive effect on house prices in a booming economy but a smaller effect in a downturn economy.

Table 7. Error-Correction Model: Short-run equation. Asymmetric effects.

	Model A5	Model B5	Model C5	Model D5
Δ Income per capita	0.9222 (5.38)	0.5230 (3.37)	0.9585 (5.61)	0.5338 (3.44)
Δ Stock per capita	-0.9818 (-4.21)	-0.8199 (-5.11)	-1.0280 (-4.43)	-0.8296 (-5.18)
Bank				
Δ Establishments	0.3458 (1.81)	-	0.3000 (1.58)	-
Δ Establishments- asymmetric	0.0861 (1.41)	-	0.0889 (1.46)	-
Δ Employments	-	0.4000 (2.90)	-	0.3953 (2.87)
Δ Employments- asymmetric	-	0.6683 (6.45)	-	0.6691 (6.47)
Δ Spatial lag (House prices)	1.0280 (12.06)	0.8714 (10.37)	1.0395 (12.22)	0.8890 (10.43)
ECT(t-1)	-0.0238 (-4.86)	-0.0188 (-3.91)	-0.0237 (-4.97)	-0.0214 (-4.55)
Constant	-0.0335	0.0059	-0.0350	0.0052

	(-3.20)	(0.50)	(-3.36)	(0.45)
AIC	-7653.21	-7765.62	-7687.23	-7769.81
BIC	-5882.77	-5988.10	-5909.71	-5992.28

Note. Dependent variable: the change of natural logarithm of house prices. We are controlling for spatial and temporal correlation as well as heteroskedasticity by estimating a feasible GLS. Time dummies are not shown in the table. t-values within brackets. The spatial matrix and the spatial lag variable have been estimated using Stata-code spwmatrix and splayvar (P.Wilner Jeanty's. 2010). AIC and BIC are used to compare the model specifications. Model D5 is preferred.

We test the hypothesis by constructing a variable that equals one for a municipality where house prices increase the following 2 years, zero otherwise. For a given municipality, the variable may thus be one for some years and zero for other years. The underlying logic is that a local bank, which has superior access to local information, may have a better “crystal ball” than non-local banks and, thus, be more accurate in foreseeing the future. Based on this reasoning, a local bank presence would be expected to have a positive impact on house prices in a booming market but a smaller - perhaps even negative - effect in a bust market. Both models that use the number of bank employees as the measure for bank presence (B5 and D5) support the hypothesis as they show a larger positive and significant, positive effect in booming markets. The results in the establishment models, A5 and C5, are positive but not significant. We do not find support for that the effect should be negative during bust periods. This result should be interpreted with care as only a small share of the observations (15%) belongs to the bust market.

Conclusions

The underlying theoretical idea of the present paper is that local banks are better informed about the local housing market than non-local banks. Possessing this information may influence the willingness of local banks to grant loans to house buyers and the specifics of the loan terms, which in turn may influence the efficiency of the housing market. Based on this reasoning, we propose a set of hypotheses that

we test these using a Swedish panel data set covering 274 municipalities over 15 years.

Our main result is that we, by using an error correction model and an instrument variable technique, manage to establish a clear causal link between local bank presence - measured as either the number of bank establishments or bank employees in a municipality - and house prices. As expected, greater bank presence, on average, yields higher house prices. The magnitude of the effect is rather stable both over time and municipality size. On average, a 1% increase in the change of the number of bank establishments (employees) per capita yields an expected increase in change in the house prices of 0.4% (0.9%).

We observe significant spillover effects; that is, bank presence in a given municipality influences house prices in nearby municipalities. We show this both using adjacent municipalities and a gravity model, however, the spatial lag representation more effectively pick up the spillover effects.

Finally, we study a “crystal ball” effect - that is, the ability of local banks to better foresee future development of the local housing market accurately. We exploit that, for a given year, we know the changes in house prices that occur in the years that follow, which, of course, was unknown to the banks. If this “crystal ball” idea is true, the local bank presence should have a larger positive effect in a booming market. In a bust market, the local bank presence should have a smaller, or even negative, impact. We find evidence for a larger positive impact in booming markets. The impact in bust markets is smaller but we do not find any support for the effect being negative.

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