THE RIMINI CENTRE FOR ECONOMIC ANALYSIS

# WP 15-39

**Mark J. Holmes**
Waikato University, New Zealand

**Jesus Otero**
Universidad del Rosario, Colombia

**Theodore Panagiotidis**
University of Macedonia, Greece
The Rimini Centre for Economic Analysis, Italy

# A PAIR-WISE ANALYSIS OF INTRA-CITY PRICE CONVERGENCE WITHIN THE PARIS HOUSING MARKET

# A pair-wise analysis of intra-city price convergence within the Paris housing market*


Mark J. Holmes†
Department of Economics
Waikato University
New Zealand

Jesús Otero‡
Facultad de Economía
Universidad del Rosario
Colombia

Theodore Panagiotidis§
Department of Economics
University of Macedonia
Greece


November 2015


**Abstract**

We examine long-run house price convergence across the twenty Paris districts using a quarterly dataset that spans from 1991 to 2014. Our econometric modelling exercise adopts a pair wise approach that is built on a probabilistic test for convergence based on house price differentials. We find that more than 50% of the intra-city house price differentials that can be computed are stationary. Our findings further reveal that the half-life of a shock to long-run price equilibrium is affected positively by unemployment, distance and housing supply.



JEL Classification: C2, C3, R1, R2, R3

Keywords: Pair-wise, house prices, cointegration, speed of adjustment.


---


*Constructive comments from an anonymous referee are gratefully acknowledge. The usual disclaimer applies.



†E-mail: holmesmj@waikato.ac.nz

‡E-mail: jesus.otero@urosario.edu.co

§E-mail: tpanag@uom.gr

# 1 Introduction

As housing constitutes the largest asset of the personal sector portfolio, understanding how regional house prices behave in relation to each other over time has stimulated considerable interest and empirical scrutiny for many years. Indeed, early work by Ball (1973) provides the first review of the literature on relative house prices evaluating six studies for the UK and five for the US. Interest in regional house price behaviour is understandable. The role of housing wealth in driving consumption expenditure motivates the attention paid by Central Banks to the state of the domestic housing market when commenting on the state of the national economy; see e.g. Case et al. (2013) for more on the housing wealth effect. But at the regional level, fluctuations in relative house prices have the potential to influence relative regional economic activity. Variations in relative prices also have the possibility to affect labour mobility (and thus unemployment) through the affordability of housing and relocation costs. In this respect, the behaviour of relative house prices across regions is important and can have implications for the necessity and form of regional adjustment policies. The literature on the interactive nexus between housing markets and the macroeconomy has been evolving rapidly and Leung (2004) provides an early review on this expanding topic.

Starting from the work of Meen, see for example Meen (1999), it has been argued that shocks to regional house prices ripple out across the economy. Whilst the notion of such a ripple effect may rely on factors such as spatial patterns in the determinants of house prices, migration, equity transfer, and spatial arbitrage, it also requires some degree of long-run constancy, or a long-run equilibrium relationship, between regional house prices. The focus of our paper is on the investigation of long-run equilibrium relationships or convergence between house prices across the districts of Paris. As noted by Abbott and De Vita (2012) in their study of house prices within London, existing studies have tended to explore how inter-regional house price diffusion operates both spatially and temporally (see Holly et al. (2011)). However, the absence of any complementary analyses examining intraregional house price convergence at this level, is rather striking. Despite the importance of Paris as a major European city, to the best of our knowledge no study has yet formally analysed house price convergence within the Parisian housing market.

While intra-city study of Paris house price convergence is absent, a limited number of researchers have applied time-series approaches to Paris house price data. For example, Roehner (1999) employs data for 20 districts of Paris (intra-muros) from 1984 to 1996. Prices in the best areas are found to peak first and decrease in proportion to their former increase. Meese and Wallace (2003) evaluate the effect of market fundamentals on housing price dynamics. Using transaction-level data for dwellings in Paris over the period 1986-92, they find evidence consistent with the hypothesis that economic fundamentals that include construction costs, the interest rate, employment and a real income proxy constrain movements in Parisian dwelling prices over longer-term horizons. Their analysis suggests that the speed of adjustment in the Paris dwelling market is about 30 per cent per month over their study period. Gil-Alana et al. (2014) estimate the fractional differencing parameter for London and Paris house price series. They find that the orders of integration are greater than one for Paris apartments signifying that the series are very persistent. In other studies of Paris house prices, Fack and Grenet (2010) investigate how housing prices react to the quality of education offered by neighbouring public and private schools. Their results confirm the predictions of general equilibrium models of school choice that private schools, by providing an advantageous outside option to parents, tend to mitigate the impact of public school performance on housing prices. Nappi-Choulet and Maury (2011) conclude that spatial and temporal drifts in household socio-economic profiles and local housing market structure effects are major determinants of the price level for the Paris housing market. Moving further afield, Vansteenkiste and Hiebert (2011) provide evidence of limited house price spillovers in the euro area. For the United States, Holmes et al. (2011) employ a pair-wise approach to investigate for the convergence of 48 US states, while Kim and Rous (2012) provide weak evidence of overall convergence for a similar dataset within a club convergence framework. Further work on US regional house prices by Miles (2015) finds substantial variation across regions and over time in terms of how integrated they tend to be. In the case of Taiwanese cities, Chien (2010) examines the issue of whether regime changes have broken down the stability of the ripple effect.

Against this background of existing studies, we contribute to the understanding of house price adjustment by examining intra-city house price convergence. In our investigation, the stationarity of house price differentials is used as an indicator of long-run regional house price convergence based on a tendency for house prices to not necessarily be equal, but instead move together over time. In this respect, our study is not about explaining price levels, but relative price differentials. As argued by DiPasquale and Wheaton (1996), one might expect house prices across all locations to rise and fall with a market's fortune, but the relative price of the more desirable versus less desirable locations perhaps changes very little in the long-run. Our analysis specifically addresses whether the expected general stability of relative prices or property price premiums is a generalised phenomenon throughout the Parisian neighbourhoods. In the spirit of the earlier studies by Holmes et al. (2011) and Abbott and De Vita (2012), we utilise an econometric procedure advocated by Pesaran (2007) and Pesaran et al. (2009) for our empirical analysis. Within this approach, a probabilistic definition of convergence is proposed and forms the basis of the test. The idea behind this is that for a sample of $N$ different Parisian neighbourhoods, which are called arrondissements, unit root tests are conducted on all $N(N-1)/2$ house price differentials. Under the null hypothesis of non-stationarity or non-convergence, one would normally expect the fraction of house price differentials for which the unit-root hypothesis is rejected to be close to the size of the underlying unit-root tests, denoted as $\alpha$. However, it can be argued that the null of non-stationarity for all state pairs can be rejected if the fraction of rejections exceeds $\alpha$. Although the underlying individual unit-root tests are not cross-sectionally independent, under the null of non-convergence (or divergence) it can be shown that the fraction of the rejections converges to $\alpha$, as $N, T \rightarrow \infty$, where $T$ is the time dimension of the panel.

In an extension to testing for long-run convergence, we also analyse the drivers of convergence. House prices represent the interaction of supply conditions and the individuals' desires to live and work in certain locales (Glaeser and Gottlieb (2009)). Regional sensitivities to demand- and supply-side factors may influence the extent of house price convergence. Factors such as labour and capital mobility may be important, but the influence on housing markets of the movement of people and firms can be complex. The usual models of spatial equilibrium argue that house prices can vary according to differences in amenities (weather, congestion, etc.) and planning rules. Regional house price interactions may occur from the gradual dissemination of information across space following any shock. In an efficient market, we might expect all regions to react at the same time to a common shock. However, there are many reasons why lags may arise in the case of housing.

In our study, we consider the variables that affect the probability of finding convergence in the Parisian house prices. If there is a shock to the relative house price between two arrondissements, then what variables will affect the speed of adjustment back to long-run equilibrium? The response to these questions enable us to contribute towards an ongoing debate addressed by earlier studies, such as Pollakowski and Ray (1997), as to whether house price relationships between contiguous states are any stronger than between non-contiguous states. This too remains an unresolved issue and we enrich the debate by considering whether distance between arrondissements is a factor that helps explain the speed of adjustment towards long-run equilibrium involving bivariate house price differentials. Further to this, we also explore the role played by demographic differences between arrondissements, differences in unemployment rates as well as differences in the growth of the housing stock.

The paper is organised as follows. The following section briefly reviews the econometric methodology that we employ. Section 3 then describes the data set and empirical analysis. Using quarterly data over a 1991q1 to 2014q3 study period, we find that evidence that is supportive of long-run convergence where the probability of convergence and speed of adjustment is significantly affected by the abovementioned drivers. Section 5 concludes.

## 2 Econometric methodology: A brief review

Our econometric modelling framework is influenced by the Pesaran (2007) pair-wise approach to analyse stochastic convergence across a large number of cross section units, which we adapt and extend to the analysis of house prices in the city of Paris. Stochastic convergence involves testing the order of integration of prices relative to a baseline (or an average) price level, and as a result the outcome of the test can be sensitive to the choice of that baseline. In contrast, the pair-wise approach that we adopt in this paper requires testing the order of integration for all possible pairs of prices and, as such, does not involve what can be a problematic choice of a single reference district in the computation of house price differentials.

In line with Pesaran (2007), we let $p_{it}$ be the observed house price series in district $i$ at time $t$, where $i = 1, ..., N$ districts and $t = 1, ..., T$ time observations. Pesaran (2007) starts off by examining the stationarity properties of all $N(N-1)/2$ possible house market price differentials (or gaps) between districts $i$ and $j$, which we denote as $p_{ijt} = p_{it} - p_{jt}$, where $i = 1, ..., N-1$ and $j = i+1, ..., N$. For this, let us consider the application of both the ADF and $\text{ADF}_{\max}$ unit root tests of Dickey and Fuller (1979) and Leybourne (1995), respectively, to the time series $p_{ijt} = p_{it} - p_{jt}$, and let us denote $z_{ij}$ as an indicator function that is equal to one if the corresponding unit-root test statistic is rejected at significance level $\alpha$ (an zero otherwise). Pesaran (2007) studies the fraction of the $N(N-1)/2$ gaps for which the unit-root hypothesis is rejected, and proposes a test statistic given by:

$$\bar{z}_{ij} = \frac{2}{N(N-1)} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} z_{ij}. \tag{1}$$

Pesaran (2007) shows that under the null hypothesis of non-stationarity (that is, divergence), the expected value of $\bar{z}_{ij}$ is equal to the chosen nominal size of the underlying unit root test, denoted $\alpha$. Within this framework, support for stochastic convergence occurs whenever $\bar{z}_{ij} > \alpha$. This setup is consistent with Definition 2.1 in Bernard and Durlauf (1995) which, applied to the housing market (in Paris), indicates that for the house prices in two arrondissements to converge in a stochastic sense, they must be cointegrated with a cointegrating vector equal to $(1, -1)'$. It ought to be noticed that this definition does not necessarily signify that in the long run the two prices are the same, but simply that they move together. In other words, the cointegrating vector serves the purpose of acting as an attractor (or long-run equilibrium relationship) such that in the short run prices may be deviate from it, but not by an ever growing amount since market conditions are expected to intervene in order to restore equilibrium. To put it another way, short-run discrepancies from equilibrium in house prices are bounded. In addition to the above definition of stochastic convergence, Bernard and Durlauf (1995) further indicate that it is also theoretically permissible to have a cointegrating vector different from $(1, -1)'$, which can be thought of as a weaker form convergence. Here, the idea behind is that the two series still move together over time, but not in the same proportion; see also Quintos (1995).

Pair-wise studies of house price convergence by Holmes et al. (2011) and Abbott and De Vita (2012) as well as pair-wise studies in other contexts such as Pesaran (2007), Nourry (2009) and Le Pen (2011) focus on computing the fraction of rejections, that is $\bar{z}_{ij}$. In what follows, we progress our investigation further by calculating not only an estimate of the proportion of bivariate relations for which cointegration is not rejected, but also by examining all individual cointegration outcomes in order to determine the drivers, if any, that help explain the findings. In addition to this, we also investigate the factors that help to explain the magnitude (in absolute value) of the price differentials, as well as those that explain the speed at which prices adjust to reach their long-run equilibrium level.

# 3 Data and empirical analysis

We employ quarterly data on the median price (in Euros) per square meter in old (ancien) apartments sold in the $N = 20$ administrative districts (arrondissements) that compose the city of Paris.[1] The price data runs from 1991q1 to 2014q3, for a total of $T = 95$ time observations for each administrative district, and concern apartments that have been sold at least twice during each quarter. The source of the price series is the database BIEN, managed by the Notary Chamber of Paris, which is available online.[2] The price series of the 20 administrative districts are plotted in Figure 1, and for the purposes of the econometric analysis are considered in logarithms. Visual inspection of the time series plots in this figure suggests that the price series move together over time and could be cointegrated across districts. Of course, for the validity of the notion that a stationary price differential implies cointegration, the underlying price series need to be non-stationary processes. The non-stationarity condition is confirmed when one applies the ADF and $\mathrm{ADF}_{\max}$ unit-root tests; these results are not reported here for brevity, but are available from the authors upon request. Another feature that is apparent from the visual inspection of the series is the existence of price differentials that in some cases reached very significant magnitudes (the largest observed difference between the maximum and minimum price in any given quarter is 150%). These two aspects

[1] There is considerably variability across the regions with the most expensive regions to be found in the west, the medium priced in the centre and in the south and lower priced in the north; see Roehner (1999) for more.

[2] The reader is referred to the Appendix for the data sources used in the paper.

are the focus of the empirical analysis that follows.

We start off by examining the stationary properties of all relative prices. Because we are interested in price differentials, rather than price levels, the use of data in nominal or real terms, the latter obtained after deflating using a national deflator, makes no difference to the results. Table 1 presents the percentage of price differentials that are stationary, $\bar{z}_{ij}$, based on all 190 differentials that can be computed using $N = 20$ administrative districts. To obtain this table, the ADF and $\text{ADF}_{\max}$ unit root tests were performed at the 5 and 10% significance levels, and the optimal lag length was chosen using the information criteria advocated by Schwarz (1978) and Ng and Perron (2001), denoted SIC and MAIC respectively, allowing for a maximum of $p_{\max} = 4$ lags. Furthermore, a trend term was included in the test regression if it was statistically significant at the 5% level. From an economic point of view, it is important to highlight that the inclusion of a time trend (if significant) can be thought of as serving the purpose of picking up effects associated to relative changes in amenities. Although at first sight it may seem odd to think of convergence as occurring when there is a linear trend in the process, let us recall that the conventional unit-root one-stage testing approach, in which the linear deterministic component is included in the test regression, is asymptotically equivalent to a two-stage approach in which the underlying time-series is first de-meaned (de-trended), and where the behaviour of the resulting series is subsequently analysed by means of a test regression with no deterministic components; see e.g. Campbell and Perron (1991). In other words, one may alternatively think that when the trend term is statistically significant, the empirical analysis is being performed using amenity-adjusted relative prices.

The unit-root test results when the optimal lag length is chosen using SIC (Table 1, top panel) indicate that at the 5% significance level both the ADF and $\text{ADF}_{\max}$ tests yield rejection frequencies of 52.6 and 55.8%, respectively. Using a 10% level, the percentages are respectively 56.8 and 61.6%. Since the rejection frequencies exceed the size of the individual ADF tests, we have evidence that the house price series across arrondissements are cointegrated with a unity coefficient. As to the results when the order of the unit-root regressions is selected using MAIC (Table 1, bottom panel), the percentage of stationary price differentials is smaller though still greater than the nominal size of the underlying unit root tests. This is because the MAIC tends to choose longer lag lengths, and this in turn reduces the power of the test. Although our subsequent findings with respect to the price differentials are qualitatively unchanged regardless of whether we use SIC or MAIC, we shall focus on the results obtained with the former criterion.

One final aspect that it is worth mentioning is related to the relevance of the trend term in the unit-root regressions. Indeed, the estimated coefficient on the trend is statistically different from zero (at the 5% significance level) in 51% of the cases (that is 98 out of 190). Incorrect omission of the trend term in the unit-root regressions yields lower rejection frequencies. For instance, at the 10% significance level the corresponding values of $\bar{z}_{ij}$ for the ADF and $\mathrm{ADF}_{\max}$ tests including only intercept are 0.263 and 0.310, respectively.[3]

Once the order of integration of the price differentials, $p_{ijt}$, is determined, we consider the pairs that are found to be stationary, and for these we compute the average price differential over the sample period, which we denote $\bar{p}_{ij} = T^{-1}\sum_t p_{ijt}$. Here, it ought to be noticed that the sub-index $t$ is dropped since we focus on stationary differentials and for these the mean and variance are constant. Table 2 reports the (absolute value of the) average price differentials that can be computed using the 20 administrative districts in which the city of Paris is divided, where the entries displayed in bold face correspond to the pairs that are found stationary based on the $\mathrm{ADF}_{\max}$ test at the 10% significance level. A closer look at this table reveals the emergence of interesting patterns. For example, if one considers the administrative districts 1, 2, 3 and 4, located in the centre of the city, evidence of stationarity is encountered in all six price differentials that can be constructed among them. By contrast, if one instead considers house prices in the more peripheral administrative districts 16, 17, 18 and 19, located in the outer border of the city, evidence in favour of stationarity is found only in the price differential that involves districts 18 and 19; for the remaining five price pairs, there is no support for the existence of long-run equilibrium relationships.

The more detailed results reported in Table 2 provide us with the motivation for further developing the Pesaran (2007) pair-wise approach in three main directions.

[3]In an additional set of estimations we also apply the Kapetanios et al. (2003) test for a unit root, against the alternative of non-linear smooth transition autoregressive (STAR) adjustment. At the 10% significance level the relative frequency of rejection of the unit-root null is 32.1%, a percentage that is slightly higher than the one obtained when using the MAIC to select the optimal number of lags

First, one might be interested in determining the drivers that affect the likelihood that $\bar{p}_{ij}$ is stationary or, to put it in another way, that $p_{it}$ and $p_{jt}$ are cointegrated with cointegrating vector equal to $(1,-1)'$. In measuring the probability that relative prices are stationary, we use the indicator function $z_{ij} = 1$ if the ADF test is rejected at the 10% significance level, and zero otherwise; to assess the robustness of the results, a similar variable is constructed using the $\text{ADF}_{\max}$ at the 10% significance level. Second, conditioning on stationary differentials, one might also wish to examine the factors that determine the magnitude of the average price differentials in absolute terms, that is $|\bar{p}_{ij}|$. Our analysis focuses mainly on the arbitrage opportunities offered by prices given by square meters, which are reflected in the magnitude of the price differentials and, rather than the sign, would require considering differentials in absolute terms. Here, the variable to be explained is given by the numbers reported in Table 2 for which $z_{ij} = 1$ (depending on the unit root test that is being used for inference). Third, conditioning again on stationary price differentials, one might alternatively be interested in finding the factors that affect the speed at which house price levels adjust when they deviate from their implied long-run equilibrium relationship. As to the speed of adjustment, for each relative price that turns out to be stationary, that is for which $z_{ij} = 1$, we compute an approximation of the half-life of a shock based on the estimated autoregressive coefficient that results from estimating an ADF-type regression.[4] To interpret the findings one must bear mind that the resulting half-life between prices in districts $i$ and$j$, which we refer to as $hl_{ij}$, is inversely related to the speed of adjustment.

Turning to the drivers that are expected to have an effect on $z_{ij}$, $|\bar{p}_{ij}|$ and $hl_{ij}$, one needs to bear in mind the need to assemble a consistent dataset for the key variables across all 20 administrative districts used in our sample. Thus, for each pair-wise $z_{ij}$, $|\bar{p}_{ij}|$ and $hl_{ij}$, the district-level drivers that that we investigate are the following.[5] First, we consider a cost or supply-side variable in the form of the average yearly percentage change in the number of housing units in district $i$ between 1990 and 2007, which we denote as $houg_i$. This variable is used to construct the differential (in absolute terms) between districts $i$ and $j$, that is $|houg_{ij}| = |houg_i - houg_j|$.

[4]The half-life of a shock is estimated with the formula $-\ln(2)/\ln(1+\hat{\delta})$, where $\hat{\delta}$ is the autoregressive coefficient in the corresponding ADF test regression; see e.g. Goldberg and Verboven (2005).

[5]Please refer to the data appendix for the sources of the data.

Second, we consider a range of demand-side variables that includes the absolute difference in unemployment rates between districts $i$ and $j$, which we denote $|u_{ij}| = |u_i - u_j|$. The inclusion of the unemployment rate is intended as a barometer of economic conditions and as an indicator of income stability (as higher unemployment indicates lower job security). Either way, this variable is expected to influence the price of houses negatively. The unemployment rate is used as an alternative to per capita income as data on the latter data would be highly problematic at this district level. In terms of demographic influences, population density can serve as a measure of demand pressure and an indirect measure of supply shortage. When the population density is high, it may imply that the land endowment is very limited and thus the possibilities to increase the supply of housing are restrained. We experimented with a general indicator of population density for each district, but only insignificant results were obtained. Instead, we incorporate demographic effects through the incorporation of the so-called old-age index calculated as the ratio between the population over the age of 65 years divided by the population between 0-14 years of age. The (logarithm of the) old-age index in district $i$ is denoted $loai_i$, so that the difference between districts $i$ and $j$ is given by $|loai_{ij}| = |loai_i - loai_j|$.[6] Age is included to capture differentiated behaviour in older households when it comes to the decision of purchasing a house in a particular area of the city in terms of mobility and speculation.

Finally, we also include the logarithm of the distance between districts, $\ln dist_{ij}$. In this case, we are particularly interested in examining whether a shorter distance is associated with a faster speed of adjustment back towards long-run equilibrium. Indeed, shorter distances between districts may facilitate arbitrage mechanisms that bring house prices into line. Following the work of Pollakowski and Ray (1997) and others, we also contribute to the debate as to whether house price relationships between contiguous states are any stronger than between non-contiguous states.[7]

In summary, the following regression models are estimated:

[6] Although it is possible to argue that the variables unemployment and/or old-age index have fluctuated over time, so that looking at one value in a specific time period is not representative, we are implicitly assuming that differences across districts have remained relatively the same.

[7] In addition to the variables listed above, we also consider other potential determinants such as population growth, and a measure of the relative strength of speculative trading versus price-supply elasticity, the latter as taken from Roehner (1999). However, all of these yield inferior results and for this reason were not included in the model specification that was finally chosen.

$$z_{ij} = \alpha_1 + \alpha_2 |u_{ij}| + \alpha_3 \ln dist_{ij} + \alpha_4 |loai_{ij}| + \alpha_5 |houg_{ij}| + u_{ij}, \tag{2}$$

$$|\bar{p}_{ij}| = \beta_1 + \beta_2 |u_{ij}| + \beta_3 \ln dist_{ij} + \beta_4 |loai_{ij}| + \beta_5 |houg_{ij}| + \varepsilon_{ij}, \tag{3}$$

$$\ln hl_{ij} = \gamma_1 + \gamma_2 |u_{ij}| + \gamma_3 \ln dist_{ij} + \gamma_4 |loai_{ij}| + \gamma_5 |houg_{ij}| + \xi_{ij}. \tag{4}$$

In terms of the signs of the coefficients, for equation (2) we expect all coefficients to be negative and significant, supporting the view that the probability of rejecting the null that relative prices are non-stationary decreases with distance, and as districts become more dissimilar in terms of unemployment rates, old age density and housing growth. By contrast, the signs for the coefficients in equations (3) and (4) are expected to be positive, suggesting that as distance increases and as districts become more dissimilar in terms of the underlying drivers, price differentials and the half-life to shocks ought to increase. Before proceeding with the presentation of the results from the cross-section estimation, it might be noted that older properties in places experiencing urban renewal and gentrification may receive more investment in terms of renovations and alterations and so on. Indeed, part of the measured price differences used in our study might potentially be driven by such investments. While data availability prevents us from including such variables as additional regressors in the models postulated above in equations (2) to (4), the inclusion of a deterministic time trend (when significant) in the earlier unit root tests reported in Table 1 goes towards incorporating such effects into our study.

Table 3 reports the results from the estimation of two probit models, one where the dependent variable $z_{ij} = 1$ if the ADF test is rejected at the 10% significance level (and zero otherwise), and the other model where $z_{ij} = 1$ if the $\text{ADF}_{\text{max}}$ at the 10% significance level (and zero otherwise). In both probit models, the estimated coefficients on $|u_{ij}|$, $\ln dist_{ij}$ and $|loai_{ij}|$ have the expected signs and are statistically different from zero, having a negative effect on the probability of finding a stationary relative price. The corresponding marginal effects are -10%, -18% and -26% respectively. The variable $|houg_{ij}|$, on the other hand, is not statistically different from zero thereby suggesting that the probability of long-run price convergence is unaffected by the relative growth in housing supply.

Table 4 reports our findings regarding the speed of adjustment of the stationary relative price series. While we would expect the speed of adjustment towards long-run equilibrium and probability of long-run convergence to be inversely related, these next set results are conditioned on those cases where cointegration is already confirmed. Here we find that the estimated coefficients on the variables $|u_{ij}|$, $\ln dist_{ij}$ and $|houg_{ij}|$ have the expected positive sign, while $|loai_{ij}|$ is not statistically different from zero. The more dissimilar are districts with respect to unemployment rates and housing growth, then the slower is the speed of adjustment towards long-run equilibrium. This time distance is of importance insofar as greater distances between districts are also associated with slower adjustment speeds.

Table 5 summarises our results regarding these same explanatory variables as potential drivers of the stationary relative price (measured in absolute value). For these models, the variables $|u_{ij}|$, $|loai_{ij}|$ and $|houg_{ij}|$ have the expected positive effect on the magnitude of relative prices. Distance, however, appears with a negative coefficient although it is not statistically different from zero, so that it can be omitted without affecting the estimated coefficients on the other variables. In the case of this variable, perhaps it is not distance per se that is important, but the transport infrastructure that is linked to distance between the administrative districts. For instance, relatively distant districts may in fact be compensated by a relatively good network between them that contributes towards a small price differential. When it comes to explaining the absolute price differences across all pairs, this might contribute towards the insignificance that we find for this variable.

Clearly, the validity of the findings of the cross-section regressions postulated in equations (2) to (4) depends upon the exogeneity of the regressors that are included in the analysis. Here, one could well argue that there may be doubts regarding the exogeneity status of $|houg_{ij}|$. Bearing in mind that this variable corresponds to the average yearly percentage change in the number of housing units in district $i$ between 1990 and 2007, we use the average rate over the years 1968 and 1990 as an instrument for $|houg_{ij}|$, and apply the Durbin and Wu-Hausman exogeneity tests. Results not reported here indicate that we are unable to reject exogeneity in the OLS equations explaining $\ln hl_{ij}$ (Table 4) and $|\bar{p}_{ij}|$ (Table 2). In the case of the probit models for $z_{ij}$ (Table 3) exogeneity could also be rejected. However, since the coefficient on $|houg_{ij}|$ is not significant, this is not likely to challenge our findings regarding the other regressors in the probit models.[8]

[8]Perhaps there are less doubts regarding the exogeneity status of the other right-hand-side

Finally, Meese and Wallace (2003) earlier suggested that the speed of adjustment in the Paris dwelling market is about 30 per cent per month. Based on our pair-wise approximations, we can compute a mean half-life of 6.2 quarters which is considerably slower. Following Nappi-Choulet and Maury (2011), we find that housing market structure may be important, but only insofar as housing growth has a positive effect on the half-life.

# 4 Concluding remarks

In this paper we have examined the long-run price convergence in the twenty districts composing Paris intra-muros. Using a dataset that runs from 1991q1 to 2014q3 for each district, a pair-wise approached has been adopted that allowed us to conduct a probabilistic test of convergence. The latter is based on the unit root testing of all pair-wise house price combinations, which is an approach that provides significant advantages over panel unit root testing procedures available in the literature. We have documented ADF rejection frequencies above 50%, such that relative prices in the districts are cointegrated with a unity coefficient. The probability of stationarity in the differential is negatively affects by unemployment differentials across districts, demographics differentials and supply-side characteristics. Last but not least, after examining the determinants of the half-life of shocks to relative prices, unemployment differentials, distance and housing stock emerge as having a positive and statistically significant effect. With regard to on-going debate concerning the strength of house price relationships between contiguous and non-contiguous regions, our analysis suggests that smaller distances between Parisian districts are associated with a faster speed of adjustment back towards long-run equilibrium.

variables. That is, distance is not expected to be affected by relative housing prices nor by their speed of adjustment. In turn, unemployment differentials are expected to depend on changing supply/demand conditions in the labour market. Lastly, the old-age index is more related to demographic transformations that change little over time.

## References


Abbott, A. and G. De Vita (2012). Pairwise convergence of district-level house prices in London. *Urban Studies 49*, 721–740.

Ball, M. J. (1973). Recent empirical work on the determinants of relative house prices. *Urban Studies 10*, 213–233.

Bernard, A. B. and S. N. Durlauf (1995). Convergence in international output. *Journal of Applied Econometrics 10*, 97–108.

Campbell, J. Y. and P. Perron (1991). Pitfalls and opportunities: What macroeconomists should know about unit roots. *NBER Macroeconomics Annual 6*, 141–201.

Case, K. E., J. M. Quigley, and R. J. Shiller (2013). Wealth effects revisited 1975-2012. *Critical Finance Review 2*, 101–128.

Cheung, Y. and K. S. Lai (1995). Lag order and critical values of the augmented Dickey-Fuller test. *Journal of Business and Economic Statistics 13*, 277–280.

Chien, M.-S. (2010). Structural breaks and the convergence of regional house prices. *The Journal of Real Estate Finance and Economics 40*, 77–88.

Dickey, D. A. and W. A. Fuller (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association 74*, 427–431.

DiPasquale, D. and W. C. Wheaton (1996). *Urban Economics and Real Estate Markets*. Upper Saddle River, NJ: Prentice Hall.

Fack, G. and J. Grenet (2010). When do better schools raise housing prices? evidence from Paris public and private schools. *Journal of Public Economics 94*, 59–77.

Gil-Alana, L. A., C. Barros, and N. Peypoch (2014). Long memory and fractional integration in the housing price series of London and Paris. *Applied Economics 46*, 3377–3388.

Glaeser, E. L. and J. D. Gottlieb (2009). The wealth of cities: Agglomeration economies and spatial equilibrium in the United States. *Journal of Economic Literature 47*, 983–1028.

Goldberg, P. and F. Verboven (2005). Market integration and convergence to the law of one price: Evidence from the European car market. *Journal of International Economics 65*, 49–73.

Holly, S., M. Hashem Pesaran, and T. Yamagata (2011). The spatial and temporal diffusion of house prices in the UK. *Journal of Urban Economics 69*, 2–23.

Holmes, M. J., J. Otero, and T. Panagiotidis (2011). Investigating regional house price convergence in the United States: Evidence from a pair-wise approach. *Economic Modelling 28*, 2369–2376.

Kapetanios, G., Y. Shin, and A. Snell (2003). Testing for a unit root in the nonlinear STAR framework. *Journal of Econometrics 112*, 359–379.

Kim, Y. S. and J. J. Rous (2012). House price convergence: Evidence from US state and metropolitan area panels. *Journal of Housing Economics 21*, 169–186.

Le Pen, Y. (2011). A pair-wise approach to output convergence between European regions. *Economic Modelling 28*, 955–964.

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

Leybourne, S. (1995). Testing for unit roots using forward and reverse Dickey-Fuller regressions. *Oxford Bulletin of Economics and Statistics 57*, 559–571.

Meen, G. (1999). Regional house prices and the ripple effect: a new interpretation. *Housing Studies 14*, 733–753.

Meese, R. and N. Wallace (2003). House price dynamics and market fundamentals: The Parisian housing market. *Urban Studies 40*, 1027–1045,.

Miles, W. (2015). Regional house price segmentation and convergence in the US: A new approach. *The Journal of Real Estate Finance and Economics 50*, 113–128.

Nappi-Choulet, I. and T.-P. Maury (2011). A spatial and temporal autoregressive local estimation for the Paris housing market. *Journal of Regional Science 51*, 732–750.

Ng, S. and P. Perron (2001). Lag length selection and the construction of unit root tests with good size and power. *Econometrica 69*, 1519–1554.

Nourry, M. (2009). Re-examining the empirical evidence for stochastic convergence of two air pollutants with a pair-wise approach. *Environmental and Resource Economics 44*, 555–570.

Otero, J. and J. Smith (2012). Response surface models for the Leybourne unit root tests and lag order dependence. *Computational Statistics 27*, 473–486.

Pesaran, M. H. (2007). A pair-wise approach to testing for output and growth convergence. *Journal of Econometrics 138*, 312–355.

Pesaran, M. H., R. P. Smith, T. Yamagata, and L. Hvozdyk (2009). Pairwise tests of purchasing power parity. *Econometric Reviews 28*, 495–521.

Pollakowski, H. O. and T. S. Ray (1997). Housing price diffusion patterns at different aggregation levels: an examination of housing market efficiency. *Journal of Housing Research 8*, 107–124.

Quintos, C. E. (1995). Sustainability of the deficit process with structural shifts. *Journal of Business and Economic Statistics 13*, 409–417.

Roehner, B. M. (1999). Spatial analysis of real estate price bubbles: Paris, 1984-1993. *Regional Science and Urban Economics 29*, 73–88.

Schwarz, G. (1978). Estimating the dimension of a model. *Annals of Statistics 6*, 461–464.

Vansteenkiste, I. and P. Hiebert (2011). Do house price developments spillover across euro area countries? evidence from a global VAR. *Journal of Housing Economics 20*, 299–314.

Figure 1: Quarterly prices (in Euros) per square meter in 20 administrative districts in Paris

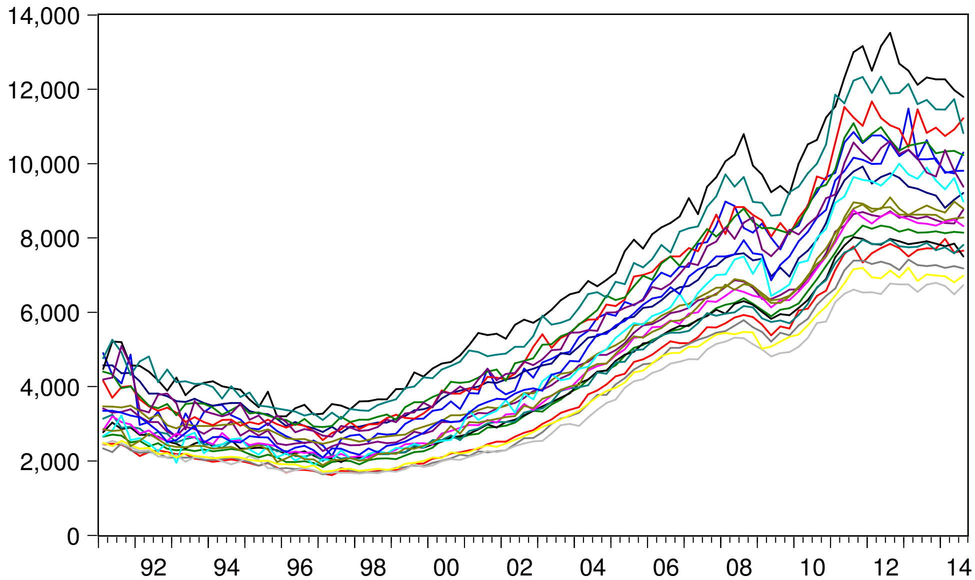

Table 1: Percentage of stationary price differentials

| Information criterion | Test | $\alpha$ | $\overline{z}_{ij}$ |
|---|---|---|---|
| SIC | ADF | 0.05 | 0.526 |
| | | 0.10 | 0.568 |
| | $\text{ADF}_{\text{max}}$ | 0.05 | 0.558 |
| | | 0.10 | 0.616 |
| MAIC | ADF | 0.05 | 0.216 |
| | | 0.10 | 0.279 |
| | $\text{ADF}_{\text{max}}$ | 0.05 | 0.216 |
| | | 0.10 | 0.295 |

*Note*: The unit root tests are performed at the significance level $\alpha$. The critical values for the ADF test are based on response surfaces estimated by Cheung and Lai (1995). For the $\text{ADF}_{\text{max}}$ test, the critical value is based on response surfaces estimated by Otero and Smith (2012).

Table 2: Average price differential between administrative districts

| Dist. | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 | 13 | 14 | 15 | 16 | 17 | 18 | 19 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 2 | **0.19** | | | | | | | | | | | | | | | | | | |
| 3 | **0.11** | **0.08** | | | | | | | | | | | | | | | | | |
| 4 | **0.05** | **0.24** | **0.16** | | | | | | | | | | | | | | | | |
| 5 | **0.07** | **0.26** | 0.18 | **0.02** | | | | | | | | | | | | | | | |
| 6 | 0.23 | **0.41** | **0.33** | **0.17** | **0.16** | | | | | | | | | | | | | | |
| 7 | **0.18** | **0.37** | **0.28** | **0.13** | **0.11** | **0.05** | | | | | | | | | | | | | |
| 8 | **0.03** | **0.22** | **0.14** | **0.02** | **0.04** | **0.20** | **0.15** | | | | | | | | | | | | |
| 9 | **0.24** | **0.05** | **0.13** | **0.29** | 0.30 | **0.46** | 0.41 | 0.26 | | | | | | | | | | | |
| 10 | **0.40** | **0.21** | **0.29** | 0.45 | 0.47 | 0.62 | 0.58 | 0.43 | **0.16** | | | | | | | | | | |
| 11 | **0.29** | **0.10** | **0.18** | **0.34** | **0.36** | 0.52 | 0.47 | 0.32 | **0.06** | **0.11** | | | | | | | | | |
| 12 | **0.27** | **0.08** | **0.16** | **0.32** | **0.34** | **0.49** | **0.45** | **0.30** | **0.03** | 0.13 | **0.02** | | | | | | | | |
| 13 | **0.27** | **0.08** | **0.16** | **0.32** | **0.34** | 0.49 | 0.44 | 0.30 | **0.03** | 0.13 | **0.02** | **0.00** | | | | | | | |
| 14 | **0.15** | **0.03** | **0.05** | **0.20** | **0.22** | 0.38 | **0.33** | **0.18** | **0.08** | 0.24 | 0.14 | **0.12** | **0.11** | | | | | | |
| 15 | **0.13** | **0.06** | **0.02** | **0.18** | **0.20** | 0.35 | **0.31** | 0.16 | **0.11** | 0.27 | 0.16 | 0.14 | **0.14** | **0.03** | | | | | |
| 16 | **0.00** | **0.19** | **0.11** | **0.05** | 0.06 | **0.22** | **0.17** | 0.02 | **0.24** | 0.40 | 0.30 | 0.27 | **0.27** | 0.16 | 0.13 | | | | |
| 17 | **0.22** | **0.03** | **0.11** | **0.27** | 0.29 | **0.44** | **0.40** | **0.25** | **0.02** | 0.18 | **0.07** | **0.05** | **0.05** | **0.07** | **0.09** | 0.22 | | | |
| 18 | **0.43** | 0.24 | **0.32** | 0.48 | 0.50 | 0.65 | 0.61 | 0.46 | 0.19 | 0.03 | **0.14** | 0.16 | 0.16 | 0.27 | 0.30 | 0.43 | 0.21 | | |
| 19 | **0.47** | 0.29 | 0.37 | **0.52** | 0.54 | 0.70 | 0.65 | 0.50 | 0.24 | **0.08** | 0.18 | 0.20 | 0.21 | 0.32 | 0.35 | 0.48 | 0.25 | **0.05** | |
| 20 | **0.42** | **0.23** | **0.31** | **0.47** | 0.49 | 0.64 | 0.60 | 0.45 | **0.18** | **0.02** | 0.13 | 0.15 | **0.15** | 0.27 | 0.29 | 0.42 | **0.20** | **0.01** | **0.05** |

*Note*: Price differentials are in absolute value. Numbers in bold face indicate that the corresponding price differential is stationary based on the $ADF_{max}$ unit root test at the 10% significance level.

Table 3: Probit models for the determinants that price differentials are stationary

| | Stationary differentials based on: | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | ADF | | | | $\text{ADF}_{\text{max}}$ | | | |
| Variable | Coeff. | (hcse) | dy/dx | (hcse) | Coeff. | (hcse) | dy/dx | (hcse) |
| Intercept | 1.998 | (0.353) | | | 2.203 | (0.352) | | |
| $\lvert u_{ij} \rvert$ | -0.320 | (0.082) | -0.095 | (0.021) | -0.298 | (0.077) | -0.085 | (0.018) |
| $ln(dist_{ij})$ | -0.615 | (0.226) | -0.183 | (0.065) | -0.673 | (0.226) | -0.191 | (0.061) |
| $\lvert loai_{ij} \rvert$ | -0.860 | (0.475) | -0.256 | (0.139) | -1.159 | (0.473) | -0.330 | (0.131) |
| $\lvert houg_{ij} \rvert$ | -0.197 | (0.561) | -0.059 | (0.166) | 0.076 | (0.567) | 0.022 | (0.161) |
| Observations | 190 | | | | 190 | | | |
| McFadden $R^2$ | 0.301 | | | | 0.313 | | | |
| LR statistic | 59.909 | [0.000] | | | 62.178 | [0.000] | | |

*Note*: Standard errors are heteroskedasticity consistent. Numbers in [•] are the probability values of the diagnostic test statistics. dy/dx denotes the average marginal effects estimated using the Delta method in Stata.

Table 4: Determinants of the half-life of stationary price differentials

| | Stationary differentials based on: | | | |
|---|---|---|---|---|
| | ADF | | $ADF_{max}$ | |
| Variable | Coeff. | (s.e.) | Coeff. | (s.e.) |
| Intercept | -1.068 | (0.196) | -0.911 | (0.217) |
| $\lvert u_{ij}\rvert$ | 0.116 | (0.052) | 0.151 | (0.059) |
| $ln(dist_{ij})$ | 0.264 | (0.132) | 0.181 | (0.145) |
| $\lvert loai_{ij}\rvert$ | 0.002 | (0.334) | -0.081 | (0.385) |
| $\lvert houg_{ij}\rvert$ | 0.863 | (0.370) | 0.971 | (0.396) |
| Observations | 96 | | 105 | |
| $R^2$ | 0.169 | | 0.144 | |
| F statistic | 4.634 | [0.002] | 4.211 | [0.003] |
| Hetero | 0.053 | [0.995] | 0.243 | [0.913] |
| Normality | 0.675 | [0.714] | 1.077 | [0.584] |

*Note*: The dependent variable is measured in logarithms. Hetero is the F-version of the White Heteroskedasticity test of unknown form, based on the auxiliary regression of the squared residuals against a constant and the squared of the original regressors. Normality is the ${\chi_2}^2$ version of the Jarque-Bera test. Numbers in [•] are the probability values of the diagnostic test statistics

Table 5: Determinants of the magnitude of stationary price differentials

| | Stationary differentials based on: | | | |
|---|---|---|---|---|
| | ADF | | $\text{ADF}_{\text{max}}$ | |
| Variable | Coeff. | (s.e.) | Coeff. | (s.e.) |
| Intercept | 0.021 | (0.025) | 0.010 | (0.025) |
| $\lvert u_{ij} \rvert$ | 0.037 | (0.007) | 0.035 | (0.007) |
| $ln(dist_{ij})$ | -0.022 | (0.017) | -0.014 | (0.016) |
| $\lvert loai_{ij} \rvert$ | 0.221 | (0.045) | 0.208 | (0.046) |
| $\lvert houg_{ij} \rvert$ | 0.204 | (0.050) | 0.273 | (0.047) |
| Observations | 108 | | 117 | |
| $R^2$ | 0.548 | | 0.552 | |
| F statistic | 31.214 | [0.000] | 34.545 | [0.000] |
| Hetero | 0.671 | [0.614] | 1.096 | [0.362] |
| Normality | 1.153 | [0.562] | 0.770 | [0.681] |

*Note*: Hetero is the F-version of the White Heteroskedasticity test of unknown form, based on the auxiliary regression of the squared residuals against a constant and the squared of the original regressors. Normality is the ${\chi_2}^2$ version of the Jarque-Bera test. Numbers in [•] are the probability values of the diagnostic test statistics

# A Data appendix

The following sources of data were consulted:

House prices:
*www.paris.notaires.fr/outil/immobilier/prix-et-nombre-de-ventes-paris-idf*

Housing units:
*www.map-france.com/Paris-75000/*

Unemployment: 2014
*www.urbistat.it/AdminStat/en/fr/classifiche/tasso-disoccupazione/comuni/paris/75/3*

Old-age index: 2014
*www.urbistat.it/AdminStat/en/fr/classifiche/indice-vecchiaia/comuni/paris/75/3*

Distance: This variable is calculated using the "greater-circle" formula based on information on latitude and longitude for the town halls in each administrative district. The geographic coordinates can be found in:
*www.map-france.com/Paris%209e%20Arrondissement-75009/*