

Cass Business School
City of London

Real Estate Finance & Investment Group

Working Papers Series

Real Estate at Cass:
http://www.cass.city.ac.uk/refig

Modelling rents in key European office markets – modern panel data techniques versus traditional approaches

Fotis Mouzakis and David Richards

REFIG WP 2004-04

Cass Business School
106 Bunhill Row
London
EC1Y 8TZ

Modelling rents in key European office markets – modern panel data techniques versus traditional approaches

Fotis Mouzakis* and David Richards**

Abstract

A significant amount of research activity covering real estate has been directed towards developing empirical models explaining rental growth - a key driver of investment performance. This paper extends the quantitative research for real estate by applying contemporary panel data techniques to 12 key European office markets. A general model is developed in which price (rent) is determined by demand and supply variables, providing a solid theoretical basis. This is tested econometrically using an error correction model which is estimated across all centres. The approach allows the behaviour of different markets to be readily compared and contrasted, providing inferences about intra-market dependence and the comparative speed of adjustment towards long run equilibrium. In addition, the model's strong basis in economic theory and parsimony help protect against misleading modelling conclusions due to the presence of poor quality data. The study extends the research by comparing the performance of the above error correction model to a number of alternatives which have been covered in the literature, including a-theoretical models. The overall results shed light on the relevance of applied economic theory for explaining real estate market behaviour, an issue which has been the subject of some debate in the literature.

Keywords: office markets, European, error correction, panel model.

* Corresponding author: Cass Business School, City University, 106 Bunhill Row, EC1Y 8TZ Tel: +44 (0)2070408212; Fax: +44 (0)2070408881; e-mail: f.mouzakis@city.ac.uk

** DTZ Research, One Curzon Street, London W1A 5PZ, England Tel: +44(0)2077576615; Fax: +44(0)2076436000; e-mail: david.richards@dtz.com

An early version of this paper was presented in the 9th conference of the European Real Estate Society, Helsinki, June 2003, under the title: "Structural variation in commercial property occupier markets - A panel data study of European office centres". The Authors are Grateful to DTZ Research for the provision of the data used in this study. The opinions expressed in this paper are those of the authors and not of DTZ

1. Introduction

A significant amount of research activity has been directed towards developing empirical models explaining rental growth - a key driver of investment performance. This paper extends the quantitative research for real estate by applying modern panel data techniques taken from applied economics to 12 major European office markets. A key benefit of this approach is that it provides new and quantifiable insights into the behaviour of European real estate markets which are of relevance to market forecasters and investment decision makers.

The econometric modelling research for real estate has predominately employed single equation methods applied to markets in isolation. This work has ranged from a-theoretical time series approaches, such as ARIMA modelling, to structural econometric modelling using a range of explanatory variables. In the latter category, D'Arcy et al [1995] and D'Arcy et al [1997] are two studies which provide interesting comparative results. On the basis of prediction accuracy tests, the authors conclude that time series methods tend to outperform structural econometric models in forecasting office markets, especially in the short term. This result could suggest that there are problems in applying economic models to real estate markets which restricts their relevance for forecasting purposes. This might occur because real estate markets do not behave in line with the way economic theory would predict or due to severe quality problems with the data, or due to a combination of both.

Another study by Chaplin [1997] draws similar conclusions. The author reviews previous modelling attempts for office rents and develops a framework for selecting models using a range of formal test criteria. The selected models are used to generate one period ahead predictions which are then compared to the out-turn. The results are not encouraging and Chaplin notes, *'With such poor performance from predictive models, it is little wonder that practitioners choose to override models' results by adjusting them subjectively to take account of expert opinion'*. In conclusion, quoting Makridakis, the author notes that, '*The use of more complex methods is one of the 'don'ts of forecasting'*.

In contrast, Mitchell and McNamara (1997), in their review of the development and application of property market forecasting, adopt a more positive stance towards the use of modelling as an aid to forecasting and investment decision making. In discussing the development of modelling for real estate in the academic literature they say, '*They [the models] tell a plausible story about what factors are important in determining rents. As such, forecasts derived from these models are likely to be credible as inputs to those making investment decisions…*'. In line with these comments, it can be inferred that the authors implicitly acknowledge the important role played by economic theory in determining which factors are plausible as explanatory variables.

Whilst the modelling work for real estate has been based largely on single equation methods for markets in isolation, it is relevant to note that panel data techniques have been developed in other areas of applied economics for the joint analysis of groups of markets taken together. It is notable that very little use of panel data analysis can be found in the real estate literature. This point is made by Hendershott et al [2002] in their panel data study which analyses UK office markets. Their analysis uses an Error Correction Model (ECM) and time series data covering demand and supply. The advantages of the ECM framework include the full parametric identification of short run and long run effects and the avoidance of spurious regression by the use of co-integrated stationary variables. The analysis yields a sensible and significant equation for offices and it is interesting that the results do not offer support for cross-sectional variation across different markets, with the exception of London. However, whilst their study employs panel data techniques, it does not follow contemporary methods developed in applied economics for testing different model types and choosing appropriate econometric methods for model estimation. For example, there is no mention of Hausman tests for the efficacy of Random Effects Models versus Fixed Effect Models.

A recent addition to the literature covering the modelling and prediction of commercial rents is Jones and Orr (2003), which provides a detailed literature review of the relevant recent research. The authors focus on three previous approaches, Wheaton et al (1997),

Tsolacos et al (1998) and Hendershott et al (1999), a predecessor of Hendershott et al (2002) reviewed above. On the grounds of these papers they propose their own model in the form of a system of simultaneous equations, hoping that stochastic interdependence between the rental equations and those for the supply side of the market will enhance the performance of the rental equations. As in Hendershott et al (2002), the econometric results based on Edinburgh and Glasgow are moderately encouraging and fall short of the performance of the model advocated in this study. Two possible reasons can be proposed for this. One is that, unlike Hendershott et al (2002), the model adopted in Jones and Orr (2003) does not include an error correction mechanism, with a restricted capability to capture the dynamics of market equilibria. Another could be that limiting a cross-sectional sample to two cities, such as Edinburgh and Glasgow, could under-represent important cross-centre influences in the urban structure, such as ripple effects from larger centres. It is worth mentioning that recent work, which models real estate markets as systems of equations, is provided by two recent conference papers, Henneberry et al (2003 and 2004). In a similar way, these theoretical efforts to develop a theory for commercial real estate markets, based on recent developments of urban theory, put more focus on the totality of the systems and less on modelling locations in isolation.

Modern panel data analysis offers a number of benefits compared to single equation methods. It allows a common demand and supply framework to be quantitatively tested across groups of real estate markets, revealing both key similarities and key differences. The testing procedures cover a wide range of relevant panel model types and this offers additional insights regarding the behaviour of real estate markets – for example, the extent to which demand and supply responses are similar or different across markets. This can be useful as an aid for isolating misleading modelling conclusions, for example, due to data problems in a particular market. Once an appropriate model has been selected, the approach involves further formal testing for selecting the appropriate econometric estimation technique, thus avoiding modelling results which are statistically inefficient, inconsistent or biased. The tests themselves can reveal further insights into market behaviour, for example, the extent to which the error terms for different markets

influence one another, a potential indicator of common but unidentified factors which are influencing prices.

This paper extends the analysis of Hendershott et al to the European arena and addresses the methodological concerns highlighted above by carefully following the full set of testing procedures required by modern panel data theory. The research also uses a dataset which has not been available in previous studies which includes DTZ prime office rents for the 12 larger European office centres, a floorspace stock index based on development completions as a proxy for supply and local market services output as a proxy for demand.

The performance of the resulting model is compared against other contemporary model specifications, which have been re-estimated using this dataset. The aim is to formally test which modelling approach offers the best performance on the basis of several criteria. The criteria selected include: goodness of fit, the predictive power for the turning point in the rent cycle which occurred in most European markets during 2001, and the aggregate size of the forecasting errors in the two year period 2000 to 2001, which was around the time of the turning point.

Section 2 discusses the specifications and the nature of the data. Section 3 deals with model testing, selection and econometric estimation. Section 4 presents the modelling results and interpretation. Section 5 compares the performance of the selected model to other contemporary model specifications. Finally, section 6 draws conclusions, considers practical applications of the model and explores how the research could be extended.

2. The data

The data set used in this study includes time series covering prime real office rents, local level GVA for market services and stylised estimates of floorspace stock. The prime rent series provides the price data for offices, the local level GVA for market services provides a surrogate for the demand for offices and the office floorspace stock series

provides a proxy for supply. The data typically runs from 1980 to 2001 and covers the following 12 locations: Amsterdam, Berlin, Brussels, Dublin, Frankfurt, London, Madrid, Milan, Paris, Rome, Stockholm and Vienna.

The prime office rent data is sourced from DTZ Research's in-house database which has been built up over time via the regular collection of prime rent estimates from DTZ's European network of local offices. The definition of prime rent varies slightly across markets according to local convention but they typically refer to centrally located units above 500sq.m. which are newly or recently constructed to a good specification. Rents are typically quoted on a net basis exclusive of service charge and local tax; the main exception is Stockholm where rents are quoted exclusive of property tax but inclusive of operating costs. To convert the prime nominal rent data into real terms, the impact of the relevant national level inflation has been stripped out. The rent data is expressed in Euro/sq.m. and is in constant 1995 prices.

The local economic data is sourced from DTZ Research/Cambridge Econometrics and typically refers to GVA for Market Services at the relevant NUTS2 region for each City. The local level GVA Market Services data is expressed in million's of Euros and is in constant 1995 prices.
For Amsterdam and Brussels, econometric analysis which has previously been conducted by the authors has not supported the use of local NUTS2 region GVA Market Services data as an explanatory variable for office rents, but has instead supported the use of national level GVA data. In view of this result, the national GVA data has been re-based to the 1995 local GVA Market Services figure for both Amsterdam and Brussels.

The stylised floorspace stock data is derived from analysing the historical relationship between floorspace stock and development completions. This relationship is used to calculate stylised stock which is based solely on the level of development completions. This has advantages for forecasting purposes as information covering future development activity is more readily available than forecasts of the floorspace stock. The floorspace

stock and development completions data is sourced from both DTZ Research and other relevant sources and is expressed in square metres.

Error Correction Models (ECM), which are based on the principle of cointegration, are introduced and tested later in this study and a discussion of the data would not be complete without some reference to the time series properties of the data. The key question is whether the series are stationary or integrated. Unfortunately, it is difficult to prove this conclusively as the relevant statistical tests are not very powerful - with only about 20 observations available for each series, the power of the Dickey-Fuller tests and the associated Augmented tests for unit roots is quite low. Despite this reservation, all but one of ADF tests failed to reject the hypothesis of unit roots, even at the 10% confidence level. The only exception was Frankfurt rents, for which the test statistic supported stationarity at the 5% level of significance. Moreover, the outcome of these tests was in agreement with the general impression from observing the plots of the time series, which broadly supports the hypothesis that the series are not mean reverting. In conclusion, the data series are assumed to be integrated of order one I(1), i.e. with stationary first differences. Whilst the ECM specification is based on the principle of co-integration, namely the estimation of the appropriate cointegrating vector for the I(1) variables, it should be noted that the relevant tests for cointegration are not very powerful, that there are a limited number of available observations, and that a number of degrees of freedom are lost to both the stationary and the non-stationary variables in the ECM. These factors effectively preclude the use of formal tests for co-integration. For this study, the relevant I(1) variables are assumed to be cointegrated – the residuals from the ECM equations are checked for broad consistency with this assumption by visual inspection. In view of the above comments, cointegration testing is not discussed further in this study.

3. Testing for model selection and estimation method

Using a single equation unrestricted Error Correction Model [ECM] has the advantage of separating long from short term effects. Annual data, as compared to those of higher

frequency, are more commonly available and less likely to suffer from problems such as timing inconsistencies and lumpiness. The strongly cyclical nature of property markets in combination with the limited length of the available series suggests a cautious approach in dynamic specification. Testing for higher order dynamics, rather than improving the estimations, can move the analysis away from fundamental economics towards the mathematical description of oscillations.

There are some critical questions that affect the development of an empirical examination based on panel data. A sensible starting point for this examination is to examine for homogeneity of the model structure over the two dimensions of the panel, i.e. over time and/or across the longitudinal dimension. Structural variation can take the form of either a variable intercept model or of a more general variable coefficient model. Heterogeneity of any of these two types may occur across either time or longitudinal (cross sectional) dimensions or across both of them. In terms of the structural variation, we have to choose between fixed and random kinds of variation. The case of fixed structural variation points to model parameters which exhibit a significant degree of stability across one or more dimensions of the panel and contrasts with random structural variation where the parameters are allowed to vary more freely; empirical results which accept the former case and reject the latter would clearly be a desirable outcome for modelling real estate markets. Altogether, these combinations of different types of heterogeneity add up to several different cases and require careful statistical testing and selection of the appropriate estimation method as well as an assessment of the compatibility with the underlying economic theory.

Another important consideration is the impact of measurement errors. Firstly, it is reasonable to expect that real estate data which is thoroughly researched rather than merely indicative will deliver more robust model estimations which are consistent with economic theory. As a consequence of this, modelling exercises can provide a useful check for assessing data quality and filtering misleading observation. This form of calibration technique can be used to asses the quality of some older observations of rental series by detecting problematic observations and investigating their validity with

reference to both experienced market practitioners and alternative data sources. Secondly, it is important to assess the nature of bias that measurement errors can incorporate into the panel data based model. As Mouzakis [1999] has shown, also quoted in Levine et al [1998], in a two variable regression with a negative and a positive slope coefficient (such as demand and supply elasticities in a price function) and the assumption that the noise in the variables is uncorrelated, the model bias due to measurement errors is downwards in absolute value when panel data techniques are used. This remains the case under the presence of an additional small size coefficient, such as the inertia of the lagged dependent variable in a simple ADRL(1) considered in this study. In other words, *the confidence for the estimated modelled impact of demand and supply on rent can be regarded as being higher than the statistical tests indicate*. This is a powerful argument in favour of applying panel data techniques to real estate market data.

For the fixed structural variation (fixed effects) model, the five relevant cases are described in detail as follows. In the case of total homogeneity over both time and cross-sections, denoted as H_1, the standard log-linear demand function for tenancy market has the form:

(1) $\Delta r_{it} = a + b_1 \Delta r_{it-1} + b_2 \Delta y_{it} + b_3 \Delta s_{it} + b_4 r_{it-1} + b_5 y_{it-1} + b_6 s_{it-1} + e_{it}$

where r_{it} is the logarithm of the rents of centre i for year t, y_{it} is the logarithm of market services ouput, i.e the proxy for local demand for prime space and s_{it} is the logarithm of the local supply side proxy, i.e. the development completions based stock index. The preliminary examination for variation of the model structure over cross-sections and time includes examination for variability of either intercepts, or all coefficients (variability of slopes with invariable intercepts is rarely a meaningful assumption). If there is evidence for variability, the question that naturally follows is about the kind of variability, i.e. if the nature of the variation is fixed or random.

As a starting point we test for variability of the intercept term. First, assuming homogeneity over time and test the hypothesis of variable intercepts across centres, H_2

(2) $\Delta r_t = a + a_i + b_1 \Delta r_{it-1} + b_2 \Delta y_{it} + b_3 \Delta s_{it} + b_4 r_{it-1} + b_5 y_{it-1} + b_6 s_{it-1} + e_{it}$,

where $\sum_i a_i = 0$ against the hypothesis of overall homogeneity H_1, using a fixed effects model estimated by OLS. Continuing, we can examine for variability of the intercept over time H_3

(3) $\Delta r_{it} = a + \gamma_t + b_1 \Delta r_{it-1} + b_2 \Delta y_{it} + b_3 \Delta s_{it} + b_4 r_{it-1} + b_5 y_{it-1} + b_6 s_{it-1} + e_{it}$,

where $\sum_i \gamma_t = 0$. Finally, we may test for the hypothesis of both variations in a two-factor fixed effects model H_4

(4) $\Delta r_{it} = a + a_i + \gamma_t + b_1 \Delta r_{it-1} + b_2 \Delta y_{it} + b_3 \Delta s_{it} + b_4 r_{it-1} + b_5 y_{it-1} + b_6 s_{it-1} + e_{it}$,

where $\sum_i a_i = \sum_t \gamma_t = 0$. If the statistical tests indicate for variation we may then investigate the kind of variation by testing for fixed against random effects models. Notice that the previous models are nested into each other in succession, i.e. H_1 is nested into both H_2 and H_3, all these are nested into H_4 whereas H_2 and H_3 are not nested into each other.

Relaxing the restriction for equal slope coefficients, we may test for variability of all coefficients over cross-sections as in H_5

(5) $\Delta r_{it} = a_i + b_{1i} \Delta r_{it-1} + b_{2i} \Delta y_{it} + b_{3i} \Delta s_{it} + b_{4i} r_{it-1} + b_{5i} y_{it-1} + b_{6i} s_{it-1} + e_{it}$,

Both the limited coverage of the cycle and the outcome of the tests reduce the scope of examining more complicated hypotheses of structural variation over time. This model has been estimated with three alternative techniques. OLS limited by the restrictive assumption of the same error variance for all sectors, Heteroscedasticity Weighted GLS allowing for block-wise heteroscedasticity (i.e. different error variance for each centre, and Seemingly Unrelated Regression allowing for a fully specified variance-covariance matrix of the errors of the centres in the sample.

For the random structural variation (random effects) model, two relevant cases are considered as follows. The time-invariant random effects model H_6 can be written as

(6) $\Delta r_{it} = a + b_1 \Delta r_{it-1} + b_2 \Delta y_{it} + b_3 \Delta s_{it} + b_5 r_{it-1} + b_5 y_{it-1} + b_6 s_{it-1} + u_i + e_{it}$

and the respective Two-Factor Random Effects model H_7

(7) $\Delta r_{it} = a + b_1 \Delta r_{it-1} + b_2 \Delta y_{it} + b_3 \Delta s_{it} + b_5 r_{it-1} + b_5 y_{it-1} + b_6 s_{it-1} + u_i + w_t + e_{it}$

where u_i and w_t are the random effects of the two cases, also assumed to follow distributions that satisfy the requirements of the classical linear regression model.

Summary statistics of these models are given in table 1 and the statistical tests between combinations of the models are given in table 2; the chosen estimation methods are described in more detail in section 4. The first 4 regressions have been estimated with OLS and the last three with FGLS. The sector invariant fixed effects model could not be estimated due to multi-collinearity, a result that needs further examination especially since it involves missing observations.

A comparison of the F-tests from the regressions clearly indicates in favour of variable coefficients over cross-sections and of variable intercepts (only) over time, with high statistical significance. The variation of the coefficients across cross-sections is clearly supported by the test of H_5 against H_2. Similarly, the variation of intercepts over time receives statistical support by the test of H_3 against H_1. While simple cross-sectional effects do not significantly improve the equation, little doubt is left about the superiority of the variable coefficient model.

Table 1: Results from the estimated fixed effects models

Model	*Type of coefficient variation*	*RSS*	*DoF*	R^2
H_1	Overall homogeneity (pooled model)	5.1043	232	0.36
H_2	Intercept over cross-sections (one factor) – Fixed Effects (FE)	4.1517	221	0.48
H_3	Intercept over time (one factor) – Fixed Effects (FE)	Multicollinearity		
H_4	Intercept over both cross-sections and time (two factors) – FE	3.0997	199	0.61
H_5	All coefficients over cross-sections (variable coefficients) – FE	1.4738	141	0.71
H_6	Random Coefficients over cross-sections (one factor)	5.3219	232	0.36
H_7	Random Coefficients – two factors	5.4470	199	0.36

Table 2: Nested tests between alternative fixed effects models

Test	*F-stat*	*D.o.F. Num.*	*D.o.F. Denom.*	*p – value*	*Conclusion*
H_2 against H_1	4.61	11	221	0.00%	Reject H_1
H_4 against H_2	3.24	21	200	0.00%	Reject H_2
H_5 against H_4	5.11	58	141	0.00%	Reject H_4

The statistics of the nested tests shown in table 1 superficially support the fixed effect models over their random effects counterparts. For completeness, table 3 presents the results of Hausman tests which formally test for the efficacy of the random effects models compared to the relevant fixed effect models. As expected, both hypotheses of random effects get strongly rejected, for either the one-factor or the two factor models. Note that the Hausman test tests a null hypothesis of identity of Fixed Effects (OLS) and Random Effects (GLS) against the alternative of biased GLS estimation. Since both estimations are used in the calculation of the test statistic this test has undesirable large sample properties under the alternative hypothesis, which namely gets accepted here.[1] However, the high statistical significance of the tests contradicts the low power of this test, enforcing the outcome that Fixed Effects is the appropriate model.

Table 3: Hausman tests between Fixed and Random Effects models

Test	χ^2 *value*	*DoF*	*p – value*	*Conclusion*
H_6 against H_2	30.45	6	0.00%	Reject H_2
H_7 against H_4	28.67	6	0.01%	Reject H_4

Table 4: LR Tests for cross-centre stochastic dependence

	RSS	Obs	Log-Likelihood	LL-ratio	Dof	chi2 p-value
OLS-WLS	1.473832	225	246.4152	9.389749	66	1.0000
SUR	1.602124	225	237.0255			

On the choice between fixed and random effects of variable intercept models, Hsiao [1986 p. 43] concludes the following: '*When inferences are going to be confined to the effects in the model, the effects are more appropriately considered fixed. When inferences will be made about a population of effects from which those in the data are considered to be a random sample, then the effects should be considered random.*' A similar logic applies in the more general case of variable coefficient model, with the difference that statistical testing for the type of variations is more complicated and doubtful. Hsiao [1996, p.136] repeats for this case that the choice between the two types depends on whether we are making inferences conditional on the individual sample or unconditional on the population characteristics.

Having strong evidence for the variation of coefficients, the question is whether we use a fixed coefficients model or a Swamy type Random Coefficients Model. According to Judge et al [1985] '*...the important consideration is likely to be whether the variable coefficients are correlated with the explanatory variables. If they are... the fixed coefficients model can be used [p.544].'* Taking this into account, from our results we have pointed out that neither variable intercept nor random effect models provide adequate explanatory power for the office markets, especially when compared with the superior result from the variable coefficients model.

The fixed effects variable coefficients model is supported by the data as the most appropriate model for the 12 European office markets under consideration. Having established the relevant model type, the next step is to consider what is the most appropriate econometric estimation technique.

For fixed effects variable coefficients model Judge et al [1985 *p.539*] clearly suggest that '*when the response coefficients... are fixed parameters... (the system) can be viewed as the Seemingly Unrelated Regression (SUR) model*'.[2] The appropriate method to estimate a heteroscedastic model depends on the state of correlation between the error terms of different cross-sections. If the disturbances of the equations of different cross-sections are

correlated, a GLS estimation of SUR gives more efficient estimates than the separate OLS estimates of each region.[3] Otherwise, the variance-covariance (V-C) matrix is diagonal, SUR reduces to the case of 'Group Heteroskedastic Model' and GLS estimation has no general efficiency gains over separate OLS regressions.[4] Note that separate OLS estimations for each market, and the identical estimates of WLS regressions have good large sample properties, providing *consistent* estimates of the V-C matrix.[5] WLS however delivered more precise statistics than OLS, whereas this was not the case with SUR model where the outlook of estimated coefficients clearly deteriorated.

The hypothesis of likelihood ratio test in table 4 is that the disturbance terms of different centres are correlated gets strongly rejected. There is more into this technical result than just a suggestion for the appropriate estimation technique, namely for either independent OLS or WLS rather than a systems WLS method. The absence of correlation amongst the error terms means that the European centres' markets are not related in any other way than via their own fundamentals. Local economies and development activity largely explain the cycle of the markets, leaving little space for other cross border influences. In other words there is evidence that that price adjustments in one centre do not travel across the European borders affecting rents in other major centres. Rental arbitrage, on a cross border basis, does not seem to have played any role in European office markets, and off-equilibrium positions do not appear to have influenced the location decisions of office occupiers. This is an interesting result that deserves further investigation, especially in view of the recent developments in European integration, with possibly important implications for the pricing of offices as an asset class and for portfolio analysis.

4. Results

Table 5 presents results from the estimation of the selected model. This shows the parameter estimates for the set of block-heteroscedastic independent equations covering each centre, estimated by Weighted Least Squares. A joint goodness of fit of 71% suggests good explanatory power, particularly for a differenced and co-integrated model.

There are no major concerns from the diagnostic tests. The Durbin-Watson statistic for the whole system of 2.24 suggests that the estimation does not have a problem with autocorrelation. A similar indication arrives from the partial auto-correlograms for each city separately where the tests for AR(1) process do not exceed the critical values from the Q-tests, whereas for higher order stay well below these levels. The Jarque-Bera test also had far from significant values in all cases except Paris, suggesting normality in the distribution of the residuals. The last two columns provide calculations of the long run coefficients which are implicit in the co-integrated vector – these are calculated by dividing the coefficients of demand and supply by the inverted coefficient of rents (i.e. $-b_{5i}/b_{4i}$). All of the long run demand and supply coefficients have the correct sign with an average of 4.4 and – 6.1 respectively. There is a moderate degree of variation across centres, indicated by the standard deviations of 1.4 and 2.5 for the demand and supply coefficients respectively. The majority of the model's slope coefficients, that is circa 81%, are statistically significant at the 10% level or higher. Note that no specific adjustments such as observation specific dummy variables have been included in the models – this is in line with study objectives, namely to test the relevance of economic theory for real estate markets using modern panel data techniques, rather than develop tools for forecasting. It is important to note that equation specific adjustments could improve the fit of the model even further - this process would need to be undertaken before the model is used for forecasting purposes

All of the short term supply parameters are highly significant. This is the most strongly supported result which delivers a clear message: *local supply plays an important role in short term rental determination*. This result could help explain the poor predictive performance, in the short term, for the structural models described in the literature which do not include a supply variable.

An interesting observation is that whilst most of the short run supply elasticities are negative, as expected, a few are also positive: Milan, Paris and Berlin. This could indicate that rents in these markets remain relatively more responsive to demand, even quite late into the rental cycle, before increased supply ultimately exerts a negative impact as the

long run equilibrium takes over. The increasing prevalence of pre-lets in the development pipeline in recent years could also be a factor – whilst pre-lets increase the supply of offices, they are effectively 'off market' in the year they complete and could still be associated with a tight market and rising rents in the short term.

Table 5: Estimates of the unrestricted ECM

Centre	Inpt^{+}	Δr_{t-1}	Δy_t	Δs_t	r_{t-1}	y_{t-1}	S_{t-1}	LR y_t^{+}	LR s_t^{+}
Amsterdam	1.25	0.10	1.65	-2.01**	-0.51**	2.35**	-2.65**	4.59	-5.18
Berlin	2.19	0.74**	8.42**	3.89**	-0.40**	1.23 *	-1.76	3.11	-4.43
Brussels	-0.19	0.34*	2.24**	-0.87**	-0.15	0.70	-0.53	4.72	-3.60
Dublin	-0.80	0.00	2.40**	-0.76**	-0.26*	0.92*	-0.97	3.57	-3.76
Frankfurt	3.34	0.36*	3.37**	-0.35**	-0.62**	1.93	-3.41	3.13	-5.53
London	3.77	-0.13	0.87	-5.40**	-0.38*	2.19**	-3.98**	5.69	-10.36
Madrid	-1.09	0.10	2.86*	-6.11**	-0.49**	3.57*	-5.13*	7.33	-10.53
Milan	-0.63	0.70**	1.62	6.77**	-0.22	0.91	-1.09	4.24	-5.07
Paris	0.73	0.25	2.37	2.53**	-0.59**	2.14**	-2.25*	3.61	-3.80
Rome	2.02	0.42*	4.11*	-4.51**	-0.58**	3.06**	-4.89**	5.26	-8.40
Stockholm	10.23	0.25	3.78**	-7.24**	-1.08**	6.55**	-10.93**	6.09	-10.15
Vienna	2.85	0.35	1.22	-3.19**	-1.07**	2.88**	-3.51**	2.68	-3.27
Average^{+}	1.8	0.3	2.5	-2.2	-0.5	2.2	-3.1	4.4	-6.1
St. deviation^{+}	2.6	0.2	1.6	4.0	0.2	1.0	1.9	1.4	2.5

* 10% one-tailed significant values, ** 2.5% one-tailed significant values

$^{+}$ Statistical significance is not indicated

All of the short term demand impacts are positive and more than half are significant. Berlin, Frankfurt, Rome and Stockholm are the fastest markets to reflect the sector's short term economic performance on market rents. London is the least responsive market with near proportionate reactions of prime rental growth to short term economic growth. However, it is important to view the short term effects alongside the long run equilibrium properties of the model. For example, London is much more sensitive to economic

growth than most markets in the long run, but the impact on rent is lagged behind the current year and this is why the short run impact is estimated as both low and insignificant.

The dynamic coefficients point out differences in the speed of rental adjustment to equilibrium across the markets, arguably providing some measure of market efficiency. According to the estimates for these coefficients, Paris and Berlin and Milan are the slowest markets to react to the latest developments of their fundamental drivers. London, Dublin and Amsterdam react the fastest, arguably demonstrating higher market efficiency.

5. Comparison of modelling performance

The analysis has demonstrated that modern panel data techniques taken from applied economics can be successfully implemented for European real estate markets. The approach provides a number of formal test statistics which offer key insights regarding the similarity and differences in behaviour across markets. Furthermore, the methodology employs additional statistical tests to ensure that the chosen model is estimated using the appropriate econometric estimation technique, thus avoiding estimations which are potentially inefficient, inconsistent or biased.

Whilst the panel based approach used in this analysis provides the above benefits, it is clearly relevant to compare the performance of the model selected by this study to some of the alternative modelling approaches which have been considered in the literature. Three alternative model specifications, all based on economic theory, were chosen for this purpose. First, a simple Autoregressive Distributed Lag model of the logarithm of rents with a first lag of the dependent variable [ARDL(1)] and the proxies of demand and supply in the right hand side of the equation, all in logs. Second, a similar model that also includes a second lag of the dependent variable [ARDL(2)], such as the model used in the RICS survey [see Key et al (1994)]– in this instance the third lag of stock, which was

used as an indicator of development activity in the RICS survey, has not been included. The reason for this omission is that this study uses development completions as a proxy for supply via the stylised floorspace stock variable and it would be therefore inappropriate to use the lagged floorspace stock variable as a predictor for development activity. Third, a fixed effects variable intercept estimation of the Error Correction Model (ECM), which is commonly implied when "panel data" methods are mentioned, such as in Hendershott et al (2002) where UK city data and stock as a proxy of supply was used. These three models were chosen for their wide spread use in applied studies and because of their individual theoretical properties. They were estimated using the same method as the selected model, i.e. weighted least squares that provide coefficient estimates which are the same as from the independent OLS regressions, but with different test statistics. In addition, a representative of the a-theoretical family of Box-Jenkins models was estimated and tested. The choice of an ARIMA(2,1,2), which implies second order auto-regressive and moving average processes on first differences, was based on the results of testing these models separately for each market. These tests indicated that second order AR and MA processes were found significant in several cases and were adopted as a common model for the entire group. In analytical terms, and using a similar notation, the three theoretically based models plus the a-theoretical ARIMA(2,1,2) can be written as:

(8) $r_{it} = a_{0i} + a_{1i}r_{it-1} + a_{2i}y_{it} + a_{3i}s_{it} + u_{it}$

(9) $r_{it} = c_{0i} + c_{1i}r_{it-1} + c_{2i}r_{it-2} + c_{3i}y_{it} + c_{4i}s_{it} + u_{it}$

(10) $\Delta r_{it} = d_{0i} + d_1\Delta r_{it-1} + d_2\Delta y_{it} + d_3\Delta s_{it} + d_4 r_{it-1} + d_5 y_{it-1} + d_6 s_{it-1} + u_{it}$

(11) $\Delta r_{it} = e_{0i} + e_{1i}\Delta r_{it-1} + e_{2i}\Delta r_{it-2} + e_{3i}u_{it-1} + e_{3i}u_{it-2} + u_{it}$

To compare model performance, three criteria are used – each of these is discussed in more detail below. In addition to the measurable tests, a discussion of the theoretical consistency of the alternative modelling approaches is also provided.

The first criterion is the goodness of fit (R^2 statistic) for whole system. Since in the two forms of ECM the dependent variable is in differences, whereas the two ARDL models are in levels, the R^2 from the estimations is not directly comparable. To overcome this problem the predicted differences were calculated from the fitted values of the two models in levels, ARDL(1) and ARDL(2). Then the total for the whole system sum of square deviations from the mean, as

$$\text{TSS} = \sum_{i=1}^{12} \sum_{t=1}^{n_i} \left(\Delta r_{it} - \frac{1}{n_i} \sum_{j=1}^{n_i} \Delta r_{ij} \right)^2 ,$$

where i indicates the city, t is the year, n_i =2001 - t_i is the number of observation for which a fitted value is available, and t_i is the oldest observation for which a fitted value is available, in most of the cases 1982. The largest number of observations was utilised by the ARDL(1) model due to its simplicity. Similarly, the total sum of squared residuals was calculated as

$$\text{RSS} = \sum_{i=1}^{12} \sum_{t=1}^{n_i} \left(\Delta u_{it} - \frac{1}{n_i} \sum_{j=1}^{n_i} \Delta u_{ij} \right)^2 ,$$

where u_{it} is the residual of city i for year t. Finally the goodness of fit for the system was calculated as $R^2 = 1 - \text{RSS/TSS}$. According to this criterion, the selected model from this study with an R^2 of 66% has a clear advantage from the other theoretically based models, with the second best being the ARDL(1) with an R^2 of 44%, around two thirds of the R^2 for the selected model. The RICS ARDL(2) model follows and the fixed effects panel has the weakest fit with an R^2 of 31%. This latter result is no surprise as the panel model restricts the slope coefficients to be equal for all cities and only allows for different intercept terms.

The second criterion is the success in predicting the turning point for 2001, the last year in the data set. In 2001 ten out of twelve cities presented negative growth after several consecutive years of positive rates, with the exception of Brussels and Milan. The test measures the number of cities that the sign of growth was predicted correctly. Of the theoretically based models, the selected model from this study predicted successfully the occurrence of a turning point in all cases. Next come the two ARDL models, both with 9

successes out of 12 each. Last, the fixed effects model produced only 5 correct predictions.

A third test measures the capacity of the model to explain rental growth during the two final years in the series, 2000 and 2001. These two years were selected because of the sharp changes, including transition from heated growth to a fall and the likelihood of disequilibrium in the tenancy markets. Before the downturn of 2001, rental growth in 2000 was unusually large in several markets, providing a good challenge for the dynamic properties of the competing models. For a measurement of the performance of the models for these two years the Mean Absolute Error (MAE) criterion was simply calculated by adding the absolute values of the residuals, always in log-differences, for the two years and the 12 cities. The MAE for the selected model from this study of 0.067 is roughly half that of those for the other theoretically based models, which are all quite close at about 0.12. The second best performer in this case is the RICS ARDL(2) model with a MAE of 0.12 followed by the fixed effects ECM panel model and the ARDL(1), but all three models show small differences in performance on this measure.

The relative merits of a number of alternative theoretically based models has been considered above. However, the fact that a-theoretical models have often outperformed parametric attempts to link commercial tenancy markets directly with economic theory has often been seen as an embarrassment to researchers. Independent of the approach one takes, such an outperformance clearly suggests that either the models used are not appropriate or the data are of very low quality for the needs of empirical studies, or both. These claims are arguably true to an extent: low frequency of the data, smoothing from aggregation and unavoidable measurement errors weaken the position of theoretically based approaches against time series methods that mainly seek for patterns of oscillations in time. For a comparison of results, various forms of ARIMA were tested separately for each city. In the majority of the cases, second order terms for auto-regressive and moving average components had significant statistical tests. Thus a pluralistic ARIMA(2,1,2) form was selected for all cities and was estimated, giving some interesting results. First, the performance of this model was clearly inferior to the performance of the selected

model from this study, in all three tests used. Secondly, ARIMA clearly outperformed the two ARDL models in two out of three tests, the overall goodness of fit and also the MAE during 2000-01, but was less successful in predicting the turning point in 2001. In other words, while the results were in line with those of previous studies, any potential benefits from time series methods vanish when a carefully selected parametric method based on economic theory is used.

Table 6 tabulates the results for each test criterion for each model specification. The 'ECM OLS-WLS Panel' refers to the model selected by the panel data approach advocated in this study. The other models, in turn, relate to the theoretically based models described by equations 8, 9 and 10 respectively above. In addition, the results for the a-theoretical ARIMA(2,1,2) model described by equation 11 is also provided.

Table 6: Predictive performance tests of the five models

Model	R-Squared (log-differences)	2000 turning point prediction success	Mean Absolute Error 2000-01
ECM OLS-WLS Panel	66.2%	100.0%	0.067
1. ARDL(-1)	44.4%	75.0%	0.127
2. ARDL(-2) – RICS	39.1%	75.0%	0.120
3. Fixed Effects Panel	31.3%	41.7%	0.123
4. ARIMA(2,1,2)	52.6%	66.7%	0.092

As a further consideration, the theoretical consistency of the predictions of the alternative models can be examined and compared to the selected model. Technically, the alternative ARDL(1) specification is a restricted form of the first order ECM selected by this study, with the short term responses to demand and supply set to be at a fixed proportion to the long term ones. Another of the alternatives, the fixed effects panel model, uses the same functional form as the selected ECM but is also a restricted version of it. It allows only limited heterogeneity across the markets, in the form of an adjustment to the average level of rental growth which is captured by the variable intercept term across the cross-sections. However, the model specification imposes a significant degree of homogeneity

by forcing the same slope coefficients across all market. In both the ARDL(1) and the fixed effects panel model specifications, econometric theory suggests that the model estimations will be biased to some unknown degree, since both specifications omit variables which have been shown to be of significant explanatory value. ARDL(1) has generally given highly significant estimates with reasonably sized speed of adjustment, always significantly different from unit roots. Also the fixed effects panel estimate of the ECM has provided a very reasonably parameterised average equation with different level adjustments for the fixed part of rental growth, i.e. the fixed effects. Due to the high restrictiveness of this model, which involves c 4.6, 3.3 and 2.8 times less coefficients than ECM, RICS and ARDL(1) respectively, it is no surprise that the predictive performance is somewhat reduced. In other words the performance of this restricted version of ECM can be seen as relatively good for its simplicity.

The ARDL(2) model has a more pluralistic form than the other models as it includes a second lag of the dependent variable, in addition to the first lag which is included by all the models. From a methodological point of view, if the statistical tests support the inclusion of the second lag dependent variable, then model selection based on the 'general to specific' technique would support its inclusion. However, a careful study of the nature of the data and models in question might suggest the opposite. The data are limited in terms of the number of observations available, they are non stationary and in several cases clearly take the shape of oscillations, due to the nature of the property cycles that most of the markets in the data set experienced during the historical period covered. The actual contribution of a second lag is to allow the model more flexibility in its adjustment to capturing the mathematical description of the historical oscillations, in a way that can capture a fairly wide range of cyclical processes. The mathematical description of cyclical movements, captured by including more than one lag of the dependent variable, can reduce, or even potentially obliterate, the contribution of the explanatory demand and supply variables prescribed by economic theory. In addition, the performance of multiple lags of the dependent variable in the estimation is interrelated with the performance of the explanatory variables covering the economic fundamentals in a rather impenetrable way. It is therefore a matter of some uncertainty as to how

statistical significance is apportioned in the model estimations between the oscillatory processes captured by the multiple lags on the one hand and economic theory which is represented by the explanatory variables on the other.

In addition to the above observations, the notion of a two year lag for the rent, alongside the one year lag, in determining today's rent, does not fit comfortably with experience of the way price is formed by the participants in the market. Most property researchers would agree that rents exhibit a degree of stickiness and do not change fully to reflect changed economic circumstances overnight - the idea that rents today take account of what rents have been in recent history seems well justified by both logic and experience. However, the assumption that the market participants also look two years back in time when setting today's rent seems surprising and is not recognised as an established market practice. In other words, the inclusion of a second year lag can mislead the model estimations and reduce performance, as it does not accord with either economic theory or market experience, even if the estimated model statistics might suggest the opposite.

From our tests it is clear that, amongst the parametric models, the simple ARDL(1) outperformed ARDL(2), giving higher goodness of fit. In our estimations, and even more markedly so in the original RICS (1994) estimations, the ARDL(2) model specification tends to give high values to the persistence coefficients, which in some cases are uncomfortably close to unit roots. This essentially further reduces the robustness of the estimates for the ARDL(2) model specification. In contrast, none of the other models gave high persistence coefficients, for any of the cities - on the contrary, they were often surprisingly low.

The higher goodness of fit reported for the ARDL(1) compared to the ARDL(2) is, at first glance, a rather surprising result as the former model is simply a restricted version of the latter. Of course, this can only happen when the goodness of fit is measured under a different equation form to that under which the models were originally estimated, i.e. when the dependent variable is measured in differences rather than in levels. If the dependent variable is in levels, the unrestricted model will report a higher R^2 by

definition. There are two possible explanations for this result. First, the potential problems associated with spurious regression could be the cause. The estimations of the stationary forms of the equation (the dependent variable measured in differences, in this instance) are better able to explain rental growth, whereas the non stationary forms (the dependent variable measured in levels) better capture the shapes of the cycles, but not the variation in rental growth. Second, the nature of prime office markets and property cycles tends to match mathematical oscillatory patterns, at least for relatively short time periods, and the patterns need not be related to the demand and supply fundamentals in the market.

The above modelling comparisons, although somewhat limited in scope, do seem to support the view that the appropriate application of economic theory can improve both the modelling results and their interpretation in an applied study of the type presented in this paper. The application of economic theory dictates both the form of the model equations and the nature of the expected stochastic processes. The former were found to be in agreement with fundamental economic rules of demand and supply. The latter pointed to independence among the alternative markets, suggesting that rental arbitrage across locations at the European level is not a significant factor in determining local level occupier demand.

It is worth noting that there are several other relevant issues which have not been discussed in this section but are nevertheless worth highlighting. For instance, the importance of the decision to estimate fixed instead of random coefficient models, based on both the outcome of Hausman tests and conceptual perceptions about the nature of variability in European prime office tenant markets, has been a key consideration for the approach advocated in this study. If the alternative random effects model had been selected, the outcome would have been very different. A further key decision was to rule out SURE estimation on the basis of the tests for the diagonal variance covariance matrix – the decision to use SURE would have provided significantly different estimates and statistics. It is both quite pleasing, and perhaps somewhat paradoxical, that the outcome

from testing several complicated models of structural heterogeneity ultimately indicated in favour of the simple independent OLS regressions.

6. Conclusions

The research demonstrates how modern panel data techniques can be successfully applied to real estate and illustrates the benefits of modelling local markets within a common demand and supply framework. The 12 European office markets all respond to supply and demand in a plausible way, as the modelled results for rental growth demonstrate - but there are significant differences between the local markets. There are different long run sensitivities to demand and supply. For example, London, Madrid and Stockholm all exhibit significantly higher long run responsiveness to both demand and supply than the other markets. There is also significant variation in the speed of adjustment to the long run equilibrium – London, Amsterdam and Madrid appear to move towards equilibrium at the fastest rate whereas Berlin, Milan and Rome are the slowest. Lastly, there is wide spectrum of short run responses to demand and supply. For example, Milan, Berlin and Paris experience positive short term supply impacts on rental growth which could indicate that rents in these markets remain more responsive to demand, even quite late into the rent cycle, before increased supply ultimately exerts a negative impact as the long run equilibrium is gradually re-established.

The estimated models capture market behaviour patterns over the historical period covered by the data and demonstrate that real estate markets behave in line with the way economic theory would predict. Furthermore, on the assumption that these past patterns are likely to persist, the models could theoretically be used to generate forecasts – this is clearly relevant to investment decision making. Given appropriate assumptions regarding future demand and supply, the models can be used to forecast rental growth. For local demand, forecasts covering market services at the European level are now available from several sources. For local supply, future development completions can be used to estimate the floorspace stock variable. However, it is important to note that the model

results have not been finely tuned for the individual markets and there is scope to improve the model estimates, for example, by using dummy variables.

In comparison to alternative contemporary model specifications, the final model selected in this study out-performs all the alternatives, on the basis of several different test criteria. Of particular relevance is the ability of the selected model to out-perform the other economic theory based alternatives during 2000 and 2001 – both in terms of predicting the turning point in the rent cycle and in terms of the lower reported modelling errors during this two year period. This favourable performance in no doubt stems from the very flexible ECM specification which has been employed. The unrestricted short run impacts of the stationary 'differenced' variables, coupled with the long run equilibrium relationships of the integrated 'levels' variables, allows the selected model to capture a much wider range of dynamic responses than the restricted alternatives. As a consequence, it is perhaps unsurprising that this model performs strongly in the comparative tests. In addition, the selected model out-performs the a-theoretical ARIMA models, a result which would be expected on the basis of the above comments. Whilst the predictive tests have been undertaken on an 'in sample' basis only, the selected model's solid foundations in economic theory and strong performance go a long way towards justifying its relevance for short term forecasting, in preference to the a-theoretical time series alternatives which have proved superior in previous studies.

It is important to note that the relevance of the estimated model as an aid for market forecasting has not been the main purpose of this study, though forecasting is clearly an important topic. Of more subtle relevance, it should be emphasised that the model results offer insights which are themselves clearly relevant for property market forecasting. An understanding of what the model results are capturing, what they are not capturing and the reasons behind this are the key issues. This distinction draws a clear line between forecasts based purely on an estimated model and a more holistic forecasting approach where the forecasting analyst interprets the model results, considers how markets might have been affected by exogenous factors and is therefore alert to any potential bias the model might encompass. As an illustrative example, Berlin exhibits a rather extreme

response to short term demand and supply with both parameters impacting positively on rental growth. However, it could be argued that some of this market behaviour is perhaps attributable to the period around German unification, a one off event, and this raises questions as to whether the estimated model parameters for Berlin are relevant for forecasting purposes. This is an important observation and highlights the role of both econometric expertise and market knowledge as pre-requisites for successful forecasting. The application of the model for forecasting purposes is an area which merits further research.

Portfolio diversification and the implementation of Modern Portfolio Theory could also potentially draw on the insights offered by the model results, particularly regarding how the different markets covary with one another. The analysis shows that there is no systematic variation in the error terms of the selected model across the markets. This is an important result which shows that covariance in the rental growth between the markets can be attributed to covariance in the supply and demand variables. This means that an analysis of the covariance of local market services and development completions, alongside the model results presented here, could yield valuable insights for portfolio diversification in European real estate markets. This is a potential area for further research.

As a further extension of the research, the methodology could be applied to other relevant groups of real estate markets. For example, the research could be extended to cover the retail sector in key markets across Europe or to include second tier European office markets. Furthermore, subject to data availability, the approach could be employed in other continents, looking at local markets across North America or the Asia Pacific region etc. At the widest geographical level, the ultimate goal could be to extend the study to examine national real estate markets within a global context – this could provide additional insights relevant to global real estate forecasting and global portfolio diversification.

References

Chaplin, R. (1999) The predictability of real office rents, *Journal of Property Research*, 16(1), pp. 21-49.

D'Arcy, E., McGough T. and Tsolacos, S. (1995) Univariate Models and Cross-Sectional Analysis of Office Rents in Twenty-Five European Cities, the 'Cutting Edge' Property Research Conference, London, September.

D'Arcy, E., McGough, T. and Tsolacos, S. (1997) An Empirical Investigation of Retail Rents in Five European Cities, *Journal of Property Valuation and Investment*, Special Issue on European Real Estate, 15(4), pp. 308-23.

Greene, W. (1993) *Econometric Analysis*, Prentice Hall.

Greene, W. (1995) *Limdep Version 7*, Econometric Software Ltd.

Hendershott, P.H., Lizieri, C. and Matysiak, G.A. (1999) The Workings of the London Office Market, *Real Estate Economics*, 27(2), pp. 365-387.

Hendershott, P., Macgregor, B. and White, M. (2002) Explaining Real Commercial Rents Using an Error Correction Model with Panel Data, Journal of Real Estate Finance and Economics, 24:1/2, pp. 59-87.

Henneberry, J., McGough, T. and Mouzakis, F. (2003a) *The Economic Impact of Planning Regimes: A Panel Data Study*, presented in the International Conference of the American Real Estate and Urban Economics Association, Cracow, Poland, June.

Henneberry, J., McGough, T. and Mouzakis, F. (2003b) *Planning regimes and their impacts on local rents and local and national output*, presented in the European Real Estate Society Conference, Helsinki, Finland, June.

Hsiao, C. (1996) Analysis of Panel Data, Cambridge University Press.

Judge, G., Griffiths, W., Carter, H.R., Lutkepohl, H. and Lee, C. (1985) *The theory and practice of econometrics*, Willey.

Key, T., MacGregor, B.D., Nanthakumaran, N. and Zarkesh, F. (1994*) Economic cycles and property cycles*, main report for Understanding the Property Cycle for the Royal Institution of Chartered Surveyors, London.

Levine, P., Mouzakis, F. and Smith, R. (1998) Prices and quantities in the arms trade, *Defence and Peace Economics*, vol. 9, pp. 223-36.

McNamara, P. and Mitchell, P. (1997) Issues in the development and application of property market forecasting: the investor's perspective, *Journal of Property Finance*, Vol. 8, No. 4, pp. 363-76.

Mouzakis, F. (1999) *Quantitative Aspects of the International Arms Trade*, Chapter of PhD Thesis, Department of Economics, University of Surrey.

Orr, M. and Jones, C. (2003) The analysis and prediction of Urban Office Rents, *Urban Studies*, Vol. 40, No. 11, pp. 2255-84.

Tsolacos, S., Keogh, G. and McGough, A.J. (1998) Modelling use investment and development in the British office market, *Environment and Planning* A, 30, pp.1409-29.

Wheaton, W., Torto, R. and Evans, P. (1997) The cyclic behaviour of the Greater London Office Market, *Journal of Real Estate Finance and Economic*s, 15, pp.77-92.

Endnotes

[1] GLS is inconsistent, see e.g. Greene [1993 p.479].

[2] A similar suggestion is made by Hsiao [1996] p.134.

[3] See Greene [1993] p.489.

[4] See Greene [1993] p.488.

[5] See Greene [1993] p.453.