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# Britain

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**Abstract.** This study examines the cyclical pattern of retail property development in Great Britain. It develops and estimates an econometric model of the volume of new development starts for retail buildings. Within the theoretical framework proposed, a dynamic specification based on changes in real retail rents and total consumer spending appears to adequately capture the cyclical variation in retail development. Changes in the values of these variables induce new retail construction within two years and an Almon polynomial lag scheme best describes the dynamic distribution of their lagged effects. Investment market influences on retail building development at the national level are not established in this study. There is also some indication of a changing economic relationship between new retail development and retail rents after mid-1995, but this can only be confirmed by appropriate tests when additional observations become available.

## Introduction

Empirical investigation of the retail building cycle should result in a better understanding of its dynamics and, more importantly, it should improve the investment and development decisions of real estate professionals who focus on retail properties. Systematic research on the determinants of the cyclical pattern of the production of retail space has only recently been undertaken, despite its importance to real estate analysts and investors. The studies of Key, MacGregor, Nanthkumaran and Zarkesh (1994a) and McGough and Tsolacos (1995, 1997) in the United Kingdom and Benjamin, Jud and Winkler (1993, 1994, 1996) in the United States have produced significant insight into the dynamics of retail property production. In particular, this work has related the cyclical movements of retail construction to the cyclical variation of prespecified variables, which are assumed to govern retail-building construction. The nature of the determining forces of retail development is likely to vary across international markets reflecting different economic characteristics, institutional environments and market structures. Benjamin et al. suggest that testing models across international markets and over different time periods is a task that is required to refine existing theories. The main objective of the present study is to model the retail building cycle in Great Britain and generate evidence on the forces and processes that underpin the cyclical pattern of retail property development at the aggregate level. Quantitative studies of retail property development are very useful to real estate professionals because the output of this work can be incorporated into the information

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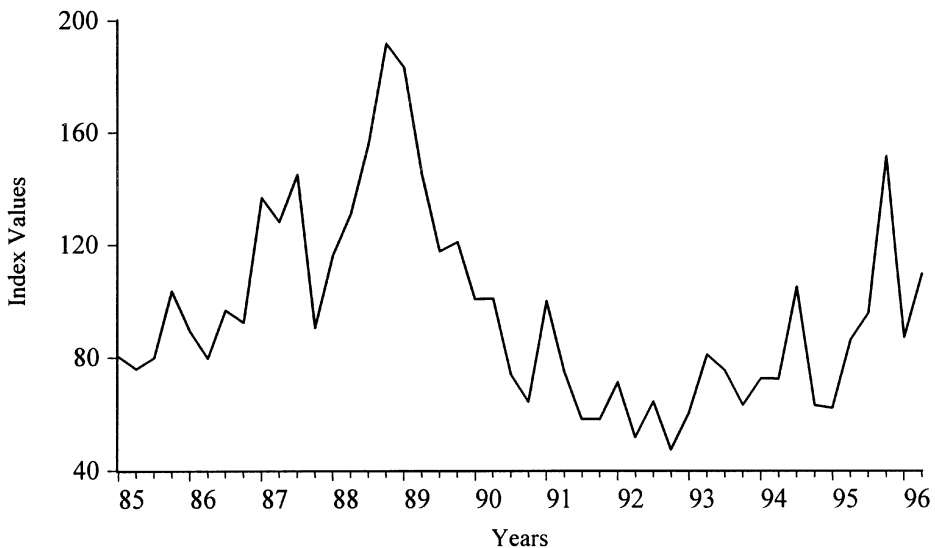
set that analysts use to form expectations about the cyclical trends of new retail development activity.

The retail building cycle in this study is defined as recurrent fluctuations in retail development and is measured by the volume of new retail orders. Therefore, the data series used in this empirical investigation of the retail cycle measures retail development starts that proxy the level of retail space completed. New retail orders relate to contracts for new retail construction work awarded to main contractors by clients in the private sector including extensions to existing contracts and speculative work undertaken on the initiative of the firm where no contract or order is awarded. Retail construction work includes shopping centers, out of town stores and small shop units. The value of this work (in sterling) is recorded in the period when foundation works are started on retail projects for eventual sale or lease. These figures are collected by the Department of the Environment in current price terms. For the purpose of this study, they are revalued at constant prices to produce a series of the volume of new retail orders in Great Britain and indexed (1990:2 = 100).

Exhibit 1 shows the retail building cycle in Great Britain, proxied by the volume of new retail orders, for the period 1985:1 to 1996:2. The series displays significant quarterly fluctuations and three different time tendencies. An upward trend is noticeable for the periods 1985:1 to 1988:4 and 1993:1 to 1996:2, whereas the period 1989:1 to 1992:4 shows a distinctive downward trend.

The empirical investigation of the forces that determine the retail building cycle in Great Britain is based on a theoretical framework within which the historical cyclical

**Exhibit 1**  
**The Retail Building Cycle in Great Britain**



variation in new retail orders is explained by: (1) changes in consumer spending; (2) conditions in the retail property market, proxied by changes in retail rents; and (3) trends in the investment market, reflected in retail property yields. The dynamic effects of these variables on new retail orders and the retail building cycle are examined with an Almon polynomial lag specification to allow for lags in the development process.

The research undertaken in this study, aiming to establish empirically the links between retail development and the variation in the a priori suggested key variables in Great Britain, contributes to our knowledge about the forces that have historically influenced the retail building cycle in the British context. The output of the econometric estimation can provide the basis to assess and predict the impact of the changing trends of the predetermined variables on the volume of new retail developments at the broader level (see also Benjamin et al. 1996). Finally, this study also provides a comparative analysis to see whether the theoretical framework, which proved successful in explaining the production of retail space in the U.S. by Benjamin et al. (1993, 1996), is appropriate to model the retail building cycle in a British context.

The remainder of this study is organized as follows. The next section reviews previous studies that provide useful insight to the modeling of retail development. The following sections present the model, methodology and results of the empirical investigation. The final section is the conclusion.

## **Related Studies**

Most of the empirical work on retail property markets in the United Kingdom focuses on rent determination and forecasting, either at the aggregate level or in more localized markets. In this literature, a number of different variables have been used as indicators of business activity in the retail sector, and by extension of demand for retail space. Hillier Parker (1984, 1985, 1987) used retail profits, disposable income and retail sales (see also Hetherington, 1988). Key, MacGregor, Nanthakumaran and Zarkesh (1994b) used consumer expenditure, but also suggested that retail sales and disposable income could be seen as alternative measures of retail activity and retail property demand. Finally, Tsolacos (1995) used the gross domestic product and consumer expenditure to estimate demand for retail space.

Other relevant research in the U.K. has focused on the determinants of retail building construction. Barras and Ferguson (1987), using spectral analysis and an error correction model, found that the main determinant of new orders for retail space in Great Britain was distributive trades output. McGough and Tsolacos (1995, 1997) examined the cyclical co-movements of the retail building cycle in relation to a number of economic, financial and property market variables, which have a priori influences on the retail sector and property market, and found that the main macroeconomic variables (gross domestic product, volume of retail sales and consumer expenditure), retail rents and capital values were procyclical and led the retail building cycle. Finally, Key et al. (1994a) modeled new retail orders, and found

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that they were influenced positively by capital values that capture profitable conditions for retail development, and negatively by interest rates and the cost of construction.

In a series of papers, Benjamin et al. have undertaken a comprehensive analysis of the retail cycle in the U.S. context (Benjamin et al. 1993, 1994, 1996). In their 1993 paper, the volume of retail sales appeared to be the main influence on the production and stock of retail space. A Koyck distributed lag model was adopted to capture the intertemporal effect of retail sales on the stock of retail space (measured by the total gross leaseable area). In this model, each successive lagged value of retail sales has relatively less weight in influencing the current supply of retail space. They found it takes over four years for the stock of retail space to increase as a result of an increase in retail sales. In their 1994 paper, they introduced the cost of capital into their models, this being measured by the real rate on long-term bonds and the tax rate, the latter proxied by the total state and local taxes as a fraction of state personal income and the average property tax rate. The cross-sectional analysis they deployed for the estimations also allowed them to examine state specific effects on retail construction. The capital and tax variables did not appear significant but state effects were found to be very important. More recently, in their 1996 paper, they used longer national time series data and found that the dominant variables in explaining new shopping center construction were retail sales lagged one period and the existing stock of shopping center retail space. Given the importance of retail sales in the modeling of retail construction, other research has focused on developing techniques to estimate retail sales including the traditional gravity model and techniques which utilize socioeconomic and demographic characteristics and shopping center specific variables (Martin, 1985; and Okoruwa, Nourse and Terza, 1994, 1996).

## Theoretical Framework

It is assumed that the cyclical variation in new retail development is induced by trends in the business of retailing, conditions in the retail property market and investors' attitudes towards retail property assets. The changing trends in the retailing sector are proxied by the variation or changes in the level of total consumer expenditure. Conditions in the retail property market are proxied by the changes in retail rents, whereas changes in retail property yields are expected to provide information about investment trends in the retail market. The functional specification between the level of new retail orders and the three variables, which underpin their variation, is given by Equation (1):

$$RENO = f(\Delta EX, \Delta RENT, \Delta RY), \quad (1)$$

where *RENO* denotes new retail orders,  $\Delta$  is the first difference operator, *EX* denotes consumer expenditure, *RENT* is retail rents and *RY* is retail property yields.

In this study, changes in total consumer expenditure, the indicator of the changing conditions in the broad economy and the business of retailing, are assumed to bring about changes in the demand for retail space. This variable has been successfully used as a demand proxy in models of retail rents in the U.K. (Key et al. 1994b; and

Tsolacos, 1995). It seems, therefore, that at the aggregate level of analysis in Great Britain, consumer spending provides an effective measure of the influence of changes in broad economic trends on the demand for retail space. Consumer expenditure is also considered by property and financial market analysts as a leading indicator both of trends in the retail sector and future general economic conditions (see *The Financial Times* 14/05/96, p. 10). In addition, economic studies have established that changes in consumer spending were the component of aggregate demand that caused the recession and recovery in the latest cycle of the British economy (Catao and Miles, 1994; Pain and Westaway 1994; and Ramaswamy, 1996). The relationship between changes in consumer spending and the volume of new retail development (new retail orders) is expected to be positive.

Changes in retail rents are assumed to reflect the state of the retail market, in particular demand-supply imbalances. Movements in retail rents also indicate the profitability of new retail projects both to developers and real estate investors and provide the signals for new retail developments. It is implied therefore that although conditions in the retail sector have an impact on the quantity of the demand for retail space, the placement of new orders and the undertaking of new developments will be subject to the current and expected demand-supply balance. Higher rental revenues should improve expectations for more profitable new retail space and encourage development starts (new retail orders). On the other hand, a weakening in rental revenues, the consequence of thin demand and surplus space, should restrain retail development activity. This means that expected rents must not only be higher, but high enough to support new construction.

It is also assumed that cyclical influences on new retail development may originate in the investment market. Investment demand for retail property is subject to wider asset market trends and is influenced by factors including the current and expected rate of return on retail investments and the risk associated with retail real estate assets. Since the current literature has not explicitly examined investment market effects on retail development, this study investigates the possibility of such influences. Changes in the level of retail property yields are included as indicators of investment demand for retail projects that are assumed to partially determine the cyclical pattern of new retail development (see also Barras and Ferguson, 1987). The relationship between new development (new retail orders) and both current and past movements in yields is expected to be negative. Falling retail property yields reflect the fact that investors favor retail property and are willing to accept a lower income return now in anticipation of higher future overall returns. As investors become more confident about this type of investment, the development of new retail projects is encouraged and the retail building cycle is reinforced.

In the functional relationship, given by Equation (1), the impacts of changes in consumer expenditure, retail rents and retail property yields are not expected to have an immediate effect on new retail development. Changes in these variables provide signals for profitable development opportunities. The time taken to assess these signals, select sites, draw up plans, obtain planning permission, secure financing, search the market to obtain tenders and so forth, will retard the response of the

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development industry and new developments will be initiated with a lag (Giussani and Tsolacos, 1994; and Benjamin et al. 1996). It could be argued that this lag tends to be longer in Britain because of the extended time needed to obtain planning permission. There are, therefore, strong reasons for expected lagged effects on new retail orders from the explanatory variables changes.

The mathematical expression of Equation (1), which provides the basis for the empirical estimation, is shown as Equation (2):

$$RENO_t = \alpha + \sum \beta_i \Delta EX_{t-i} + \sum \gamma_j \Delta RENT_{t-j} + \sum w_h \Delta RY_{t-h} + \varepsilon_t \quad (2)$$

for  $i = 0, 1..I$      $j = 0, 1..J$      $h = 0, 1..H$ .

In Equation (2) the variables  $RENO$ ,  $\Delta EX$ ,  $\Delta RENT$  and  $\Delta RY$  are defined as in Equation (1).  $t - i$ ,  $t - j$  and  $t - h$  denotes the lagged effects of the three right hand side variables on new retail orders. The maximum lag lengths are given by  $I$ ,  $J$  and  $H$  quarters. The term  $\alpha$  is the constant and  $\varepsilon_t$  is the stochastic term. The coefficients of the variables  $\Delta EX_{t-i}$ ,  $\sum \gamma_j \Delta RENT_{t-j}$  and  $\Delta RY_{t-h}$  are given by  $\beta_i$ ,  $\gamma_j$  and  $w_h$ . The terms  $\sum \beta_i \Delta EX_{t-i}$ ,  $\sum \gamma_j \Delta RENT_{t-j}$  and  $\sum w_h \Delta RY_{t-h}$  represent the sum of the contemporaneous and lagged values in  $\Delta EX$ ,  $\Delta RENT$  and  $\Delta RY$  on new retail development. There are no a priori restrictions on the maximum lag length or on the lags that are most important. These will be determined by the estimates. Benjamin et al. (1993) have argued that the current level of retail stock is the result of changes in the volume of retail sales approximately four years in the past. In the estimates of the present study, it is expected that a lag of up to two years is sufficient to capture the impacts of the independent variables on the current level of new retail orders. However, if the results indicate that longer lags are significant they will be included in the model. The data used for all variables are national series and are adjusted for inflation. Full definitions of the data and data sources are given in the Appendix.

The expected dynamic lagged effects of  $\Delta RENT$ ,  $\Delta EX$  and  $\Delta RY$  on  $RENO$  can be examined by imposing theoretical restrictions on the coefficients of the lagged values of these variables. The Koyck distributed lag model, for example, used in the study of retail production by Benjamin et al. (1993, 1996), assumes that the coefficients of the lagged values of the independent variables are declining continuously following the pattern of a geometric progression. The weight, therefore, of each lagged value of the independent variables is a fraction of the weight of the previous period. This study uses an Almon distributed lag structure, which follows an inverted V pattern. According to this pattern, the effect of the most recent changes in consumer expenditure, retail rents and retail yields on new retail orders may be negligible because of all the economic and institutional factors involved in the decision making before new developments are initiated. The main effect of the changes in these variables may be experienced in the market after a few quarters and thereafter this effect might decline again. Therefore, while coefficients from a simple Koyck model are constrained to a steadily decreasing pattern, the Almon lag model can allow coefficients to first increase and then decrease. This appears to be the pattern that is actually manifested in the market for new retail development.

The estimation period for Equation (2) is constrained by the availability of the data for new retail orders. The start date for this series is 1985:1. The latest available observation is in 1996:2 at the time of conducting the empirical investigation. Therefore, the present study makes use of quarterly data and the sample period is 1985:1 to 1996:2. However, data for the explanatory variables are available prior to 1985:1 (the quarterly series for retail rents and retail yields begin in 1977:2, whereas the data for total consumer expenditure are available from 1955:1. Therefore, when the lagged values of these variables are used, degrees of freedom are not lost and the empirical investigation is carried out with 46 observations (1985:1–1996:2).

## Methodology

The empirical investigation of the retail building cycle in the British context begins with the estimation of a Koyck model of new retail orders. The Koyck framework, as suggested by Benjamin, et al. (1993), is based upon two theoretical premises. First, the total gross leaseable area of shopping space in an area (*TGLA*) is a function of the past volume of retail sales (*RS*) as Equation (3) illustrates:

$$TGLA_t = f(RS_{t-1}, RS_{t-2}, RS_{t-3}, \dots). \quad (3)$$

Second, the effect of retail sales on the supply of retail space is captured by the Koyck distributed lag structure. The Koyck structure can be transformed into a simple equation (see Benjamin et al. 1996) that economizes on degrees of freedom and tends to avoid the multicollinearity problems that arise when several lagged values of the explanatory variables are used in an econometric model. Benjamin et al. (1993, 1996) transformed the initial Koyck distributed lag structure of retail space production and derived an autoregressive equation for the supply of retail space, which included the level of retail sales and the stock of retail space both lagged one period. The model these authors estimated is given by Equation (4):

$$TGLA_t = \alpha_0 + \alpha_1 RS_{t-1} + \alpha_2 TGLA_{t-1} + e_t, \quad (4)$$

where  $\alpha_0$  is the constant and  $e_t$  the stochastic error term.

Equation (4) produced satisfactory results in modeling the level of the total gross leaseable shopping center space in the U.S. markets considered. The authors argued that past sales have strong impacts on the current levels of shopping center space. Equation (4) provides the basis for the Koyck model of the retail building cycle in Great Britain expressed in Equation (5). The retail sales data series used in Equation (5) is described in the Appendix.

$$RENO_t = \beta_0 + \beta_1 RS_{t-1} + \beta_2 RENO_{t-1} + u_t. \quad (5)$$

The next step in the analysis is to estimate a model of new retail orders using Equation (2). Initially the correctness of the assumptions made about the causal structure of the model should be examined. That is, whether the variation in new retail orders is indeed caused by changes in consumer spending, retail rents and retail property yields.

For this purpose, Granger (1969) causality tests are performed for each independent variable (within a bivariate VAR system).<sup>1</sup> The Granger test involves testing jointly for the significance of the lags of the relevant explanatory variables ( $\Delta EX$ ,  $\Delta RENT$  and  $\Delta RY$ ). Thus, in testing the causal relationship between *RENO* and  $\Delta EX$ , we will say that  $\Delta EX$  Granger causes the variation in *RENO* if, in the equation of *RENO*, the null hypothesis of zero lagged coefficients of  $\Delta EX$  is rejected, while in the equation of  $\Delta EX$  the null hypothesis of zero coefficients of *RENO* is not rejected. Similar hypotheses will be tested for *RENO* and  $\Delta RENT$  and *RENO* and  $\Delta RY$  (see Endnote 1). Therefore, the procedure of carrying out Granger causality tests involves the estimation of regressions, which include lags of both the dependent and the independent variables. The number of lags is usually taken arbitrarily but it is advisable to run tests for different lags in order to examine the sensitivity of the results to these lag lengths. When quarterly data are used, as in the present study, it is customary to run tests for four and eight quarters.

The final statistical model that provides the most appropriate framework for the study of new retail orders and the retail building cycle in Great Britain will be derived from the empirical estimation of Equation (2) over the sample period. In estimating Equation (2), the assumption made about the lagged effects of the independent variables on the current level of new retail orders, that due to the lagged response of retail developers to demand signals the coefficients on the independent variables are expected to follow an inverted V pattern, should be examined. In order to test whether the lagged effects of the independent variables on *RENO* are captured by this pattern, an Almon polynomial lag procedure is used.

In general, the Almon polynomial distributed lag model is an estimation procedure for distributed lags that allows the coefficients of the lagged independent variables to follow a variety of patterns as the length of the lags increases (Almon, 1965). The degree of these polynomials may vary in order to capture the different lag distributions that are suggested by economic theory. Since an inverted V pattern is assumed in the present study, a turning point is expected and, therefore, a second-degree polynomial is considered appropriate to characterize this lag structure. In addition, the use of Almon lags requires the determination of the maximum lag length, which usually is a question of judgement. One way to choose the lag length is to maximize the adjusted  $R^2$ . In the present investigation, the maximum lag length is assumed to be two years (eight quarters). However, in estimating the pattern of the lagged impacts of the independent variables on new retail orders using Almon lags, end restrictions on the coefficients are not imposed. If the results suggest that the lag length is likely to be longer, that is the adjusted  $R^2$  is maximized when the lag length is over eight quarters, additional lags will be included and their statistical significance examined.

The estimation of Equation (2) follows the general to specific approach. The maximum length of the lags of the independent variables is that identified by the Almon polynomial estimations. The final model will contain only the lags of the independent variables that embody statistically significant information to explain the variation in new retail orders in Great Britain. Variable deletion tests are used to identify and omit the lags that did not contain significant information.<sup>2</sup> The robustness of the final model



is examined with a number of diagnostics tests. The Breusch-Godfrey test (Breusch, 1978; and Godfrey, 1978) for higher order of serial correlation, the White test (White, 1980) for heteroskedasticity and Wald tests (Wald, 1943) for omitted variables are applied. These tests will establish whether persistent influences on new retail orders originating in variables omitted from the final equation are not considered, and whether the final model is correctly specified. In addition, particular emphasis is given to the existence of structural breaks over the estimation period. This is important because structural breaks are associated with time variant coefficients and with a changing relationship between the independent variables and new retail orders between two different periods.

## Empirical Results

The results of replicating the model of the retail building cycle proposed by Benjamin et al. (1993) in Great Britain are reported in Exhibit 2. Although new retail orders lagged one period (quarter) appear significant, the coefficient of retail sales is insignificant and takes the negative sign that is not expected a priori. The model explains 52% of the variation in new retail orders and this can be entirely attributed to its previous levels. However, given that the retail sales variable is not significant and does not take the expected sign, it appears that this specification does not provide much insight into the forces that have historically determined the cycle of new retail orders. Therefore, an issue that needs further investigation is the extent to which the independent variable retail sales adequately captures demand side influences in the retail sector.

The different results obtained from the estimation of the Koyck model in the British context may reflect particular dissimilarities in the way the development industry assimilates and analyzes information in the U.S. and British retail markets. Retail property developers and investors in the U.S. may consider retail sales as a better indicator in the analysis of retail business trends. Moreover, the results may be partially subject to the different estimation procedures and sample periods used in the

**Exhibit 2**  
**Estimation of New Retail**  
**Development Using the Koyck**  
**Distributed Lag Model**

Variable	Coefficient	t-ratio
<i>CONSTANT</i>	49.53	1.0
<i>RS<sub>t-1</sub></i>	-0.23	0.5
<i>RENO<sub>t-1</sub></i>	0.73	6.9

Notes: The dependent variable is *RENO<sub>t</sub>*; adjusted  $R^2 = 0.52$ ; *DW h*-Statistic = -1.15. The sample period is 1985:2-1996:2 and the number of observations is 45.

study of Benjamin et al. (1993) and the present investigation. The former study estimated a cross-sectional model and the sample data included observations from 100 counties in North Carolina for 1987 and 1988. The latter study estimates a model using national time series data and the sample period is from 1985:1 to 1996:2. On the other hand, the assumptions of the Koyck model may not provide the most appropriate structure for the study of the dynamic behavior of the relationship between demand and supply in retail property development in Great Britain.

Before utilizing the framework described by Equation (2) to examine new retail development in Great Britain, the causal relationships assumed in this framework are examined using Granger causality tests. The results are shown in Exhibit 3. With regard to the effect of rent changes ( $\Delta RENT$ ) on new retail orders ( $RENO$ ), the null hypothesis  $H_0$ , that  $\Delta RENT$  does not cause variation in  $RENO$ , is strongly rejected in both cases when the lag length (denoted by  $m$ ) is four quarters ( $m = 4$ ) or eight quarters ( $m = 8$ ) (since the computed  $F$ -Statistic ( $F = 6.41$ ) is higher than the  $F$ -critical value at the 5% level of significance and appropriate degrees of freedom ( $F(4,33) = 2.66$ )). On the other hand, the null hypothesis  $H_1$  that  $RENO$  does not cause  $\Delta RENT$  cannot be rejected in either case (observed  $F$ -Statistic is less than the critical value). Therefore, based on these tests, it is clear that changes in real retail rents cause variation in the volume of new retail orders. The same conclusion cannot

**Exhibit 3**  
**Granger Causality Tests for the**  
**Independent Variables of Equation (2)**

Null Hypothesis	$F$ -Statistic
Panel A: Number of Lags: $m = 4$	
$H_0$ : $\Delta RENT$ does not cause $RENO$	6.41
$H_1$ : $RENO$ does not cause $\Delta RENT$	1.06
$H_0$ : $\Delta EX$ does not cause $RENO$	4.39
$H_1$ : $RENO$ does not cause $\Delta EX$	1.63
$H_0$ : $\Delta RY$ does not cause $RENO$	0.28
$H_1$ : $RENO$ does not cause $\Delta RY$	1.43
Panel B: Number of Lags: $m = 8$	
$H_0$ : $\Delta RENT$ does not cause $RENO$	4.02
$H_1$ : $RENO$ does not cause $\Delta RENT$	0.97
$H_0$ : $\Delta EX$ does not cause $RENO$	3.06
$H_1$ : $RENO$ does not cause $\Delta EX$	4.21
$H_0$ : $\Delta RY$ does not cause $RENO$	0.81
$H_1$ : $RENO$ does not cause $\Delta RY$	1.39

Note: In Panel A the critical value  $F(4,33)$  at 5% = 2.66; the sample period is 1986:1–1996:2; the number of observations is 42. In Panel B the critical value  $F(8,21)$  at 5% = 2.42; the sample period is 1987:1–1996:2; the number of observations is 38.

be reached for the effect of changes in retail yields ( $\Delta RY$ ) on  $RENO$ . In both cases, the hypothesis that  $\Delta RY$  does not cause variations in  $RENO$  cannot be rejected. Similarly, the hypothesis that  $RENO$  does not cause  $\Delta RY$  cannot be rejected. These tests, therefore, suggest that no relationship exists between  $RENO$  and  $\Delta RY$ . Possible reasons could be that changes in retail yields is not a suitable variable to capture the effects of the investment market on retail development or that the influence of the investment market on the retail building cycle cannot be established at the aggregate level of analysis. Investment influences may, however, be of relevance in certain geographical markets and particular types of retail property. The use of levels of yields, instead of changes, did not produce any better results.

The results obtained for changes in consumer expenditure ( $\Delta EX$ ) indicated a clear rejection of the null hypothesis that  $\Delta EX$  does not cause movements in  $RENO$  both when both  $m = 4$  and  $m = 8$ . However, the hypothesis that  $RENO$  does not cause  $\Delta EX$  is not rejected when  $m = 4$  (observed  $F$ -test = 1.63; critical value  $F_{0.05} = 2.66$ ), but it is rejected when  $m = 8$  ( $F$ -test = 4.21; critical value  $F_{0.05} = 2.42$ ). Thus, when eight lags are included in the estimations it appears that  $RENO$  also causes  $\Delta EX$ . The effect of  $RENO$  on  $\Delta EX$ , when more remote lags are considered in the Granger causality tests, could be the outcome of interactions between the wider economy and retail property market as the time horizon lengthens. Nevertheless, the overall results clearly establish the effect of  $\Delta EX$  on  $RENO$ .

Based on these findings, it was therefore considered appropriate to modify Equation (2) for the empirical estimation by excluding changes in real retail yields. The model of new retail orders can now be written as:

$$RENO_t = \alpha + \sum \beta_i \Delta EX_{t-i} + \sum \gamma_j \Delta RENT_{t-j} + \varepsilon_t \quad (6)$$

for  $i = 0, 1..I$       $j = 0, 1..J$ .

The pattern of the lagged impacts of  $\Delta EX$  and  $\Delta RENT$  on  $RENO$  is examined with the introduction of Almon lags in the estimations. The results obtained from the unrestricted and Almon equations appear in Exhibit 4. The effects both of contemporaneous and lagged values of  $\Delta RENT$  and  $\Delta EX$  on  $RENO$  are shown based on ordinary least squares estimates of: (1) the unrestricted distributed lag equation where the coefficients are estimated directly; and (2) a second degree polynomial lag equation where the coefficients are constrained to lie on a polynomial of second degree. Following the assumption that a lag length of eight quarters corresponds to the expectation about the longest time over which each of the independent variables  $\Delta RENT$  and  $\Delta EX$  could help predict  $RENO$ , eight lags were initially included in the estimations. However, the actual lag length was determined by maximizing the adjusted  $R$ -squared in the Almon equations. This appeared to be seven lags for both  $\Delta EX$  and  $\Delta RENT$ . The results for  $\Delta RENT$ , based on the unrestricted equation, show that the estimated coefficients, with the exception of those for  $\Delta RENT$  and  $\Delta RENT_{t-1}$ , appear to follow an inverted V pattern. However, the  $t$ -ratios establish the significance of the third lag only. Although most of the coefficients fail to pass the  $t$ -tests, the equation  $F$ -Statistic establishes the joint significance of this group of  $\Delta RENT$  lags.

**Exhibit 4**  
**Unrestricted and Almon Estimates of the Lagged Distribution Patterns**

Variable	OLS Estimation		OLS Estimation	
	Unrestricted Distributed Lag Equation		2nd order Almon Lag Equation	
	Coefficient	t-ratio	Coefficient	t-ratio
<b>Panel A: <math>\Delta RENT</math></b>				
$\Delta RENT$	1.60	0.6	-0.86	0.6
$\Delta RENT(-1)$	-2.25	0.7	1.62	3.1
$\Delta RENT(-2)$	2.75	0.9	3.22	5.5
$\Delta RENT(-3)$	7.17	2.3	<b>3.94</b>	4.4
$\Delta RENT(-4)$	3.29	1.0	3.77	4.2
$\Delta RENT(-5)$	0.84	0.3	2.71	4.6
$\Delta RENT(-6)$	2.60	0.8	0.77	1.4
$\Delta RENT(-7)$	-2.85	1.1	-2.06	1.4
Adjusted $R^2$		0.66		0.68
DW Statistic		1.52		1.65
F-Statistic		12.02		32.59
<b>Panel B: <math>\Delta EX</math></b>				
$\Delta EX$	3.40	0.6	-3.02	0.8
$\Delta EX(-1)$	-2.07	0.4	2.19	1.3
$\Delta EX(-2)$	0.75	0.1	6.04	3.8
$\Delta EX(-3)$	3.96	0.9	8.52	3.9
$\Delta EX(-4)$	19.74	3.4	<b>9.63</b>	4.4
$\Delta EX(-5)$	11.10	1.9	9.38	5.8
$\Delta EX(-6)$	5.82	1.1	7.76	4.6
$\Delta EX(-7)$	2.63	0.5	4.77	1.2
Adjusted $R^2$		0.60		0.60
DW Statistic		1.27		1.36
F-Statistic		9.45		23.69

Note: The sample period is 1985:1–1996:2 and the number of observations is 46.

This points to strong collinearity between lagged values of  $\Delta RENT$ , a problem which is not uncommon in unrestricted distributed lag models. However, as expected, the use of Almon polynomials to test for the significance of the inverted V pattern remedies this problem. The results show that the computed values of the constrained coefficients of the lags of  $\Delta RENT$  are significant. The turning point occurs after three lags. The  $t$ -ratios also show that changes in real retail rents help predict the current level of new retail orders when lagged between one and five quarters. The adjusted  $R^2$  has improved slightly and the value of the  $DW$ -Statistic is satisfactory.

The results for  $\Delta EX$  are not dissimilar. Multicollinearity problems are detected when the unrestricted distributed lag equation is estimated (high adjusted  $R^2$  and insignificant  $t$ -ratios). Again, with the exception of the first two coefficients, an

inverted V lag distribution pattern emerges. This pattern becomes more evident when the use of a second-degree polynomial corrects for multicollinearity in the unrestricted estimation. The initial significant effect is established after two lags (quarters) and the last one after six lags. The results, therefore, suggest that  $\Delta EX$  lagged by two to six quarters captures the full effect of changes in expenditure on new retail orders. The peak occurs at lag  $t - 4$ . The *DW*-Statistic, although higher when polynomial lags are used, still points to a possible positive first order serial correlation. The use of an AR(1) iterative procedure, instead of OLS (for both the unrestricted and Almon lag equations) to correct this problem improves the *t*-Statistics, but the results do not change. The coefficient on the fourth lag of  $\Delta EX$  was the largest and the range of the significant lagged effects was again found to be between two and six quarters. It can also be observed in Exhibit 4 that the adjusted  $R^2$  does not increase when the Almon polynomial is used. This may suggest that the use of the Almon polynomial is not necessary to impose restrictions on the coefficients to follow the inverted V pattern, since the coefficients on  $\Delta RENT$  and  $\Delta EX$  in the unrestricted equation capture this pattern to a degree. However, the multicollinearity problems associated with the unrestricted equation necessitate the use of Almon lags to establish the statistically significant coefficients of different lags and the dynamic pattern of influences of  $\Delta RENT$  and  $\Delta EX$  on new retail orders.

Equation (6) is now estimated in its general form with seven lags of  $\Delta RENT$  and  $\Delta EX$  (as identified by the Almon polynomial lag procedure) and restrictions on the coefficients to follow the inverted V pattern. The simultaneous inclusion of seven lags of each independent variable resulted in collinearity problems between the regressors (despite the use of Almon polynomial lags for each of the independent variables) and several lags of  $\Delta RENT$  and  $\Delta EX$  lost their statistical significance. The variable deletion tests (outlined in Endnote 2) were used to omit the lags, which did not contain significant information in explaining the variation in new retail orders. The application of these tests resulted in a specific model of new retail orders which is given as Equation (7). Equation (7) suggests that the effect of  $\Delta RENT$  and  $\Delta EX$  on new retail orders in the sample period 1985:1 to 1996:2 can best be captured by the third lag of  $\Delta RENT$  and the fourth lag of  $\Delta EX$ . This is the preferred dynamic specification for modeling the cyclical movements of retail development in Great Britain.

$$RENO_t = 81.30 + 9.02\Delta RENT_{t-3} + 14.82\Delta EX_{t-4}. \quad (7)$$

Exhibit 5 reports the output of diagnostics tests for this equation. Equation (7) explains 73% of the variation in new retail orders. This is good explanatory power, given the volatility exhibited in this series. The Breusch-Godfrey test for serial correlation, which tests for the presence of more general forms of autocorrelation, takes a value of 6.42 which is less than the critical value of  $\chi^2(4) = 9.49$  at the 5% level of significance. That is, serial correlation, up to fourth order, is not significant in Equation (7). Similarly, the White test for heteroskedasticity, which also follows the  $\chi^2$  distribution, takes the value of 0.68, which is much lower than the critical value of  $\chi^2$  with five degrees of freedom ( $\chi^2(5) = 11.10$ ). Therefore, the OLS estimates are efficient, the error variances constant and they remain the same irrespective of small or large values of the explanatory variables. The influence of a time trend on the

**Exhibit 5**  
**Empirical Estimation of the Retail Building Cycle**

Variable	Coefficient	t-ratio	Probability
<i>CONSTANT</i>	81.30	23.3	0.00
$\Delta RENT_{t-3}$	9.02	5.8	0.00
$\Delta EX_{t-4}$	14.82	3.3	0.00

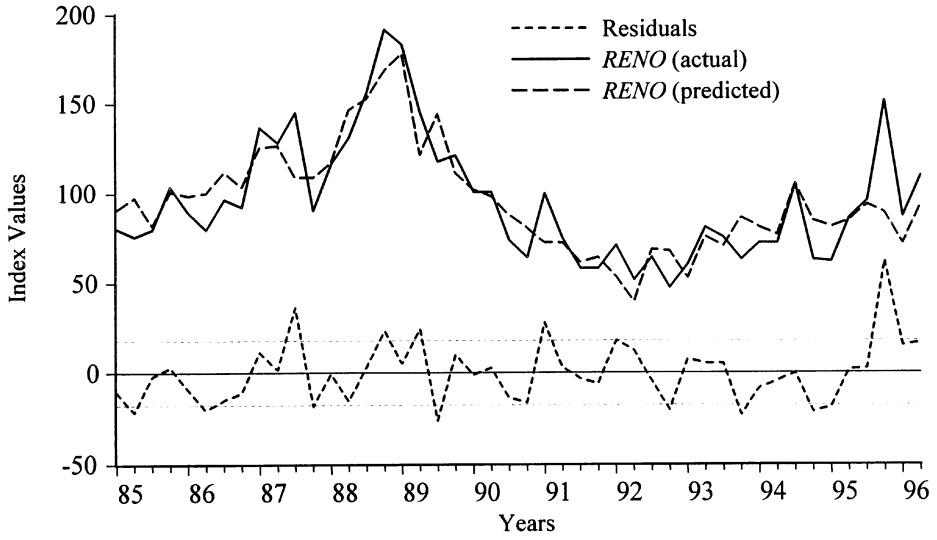
Note: This exhibit is based on Equation (7). The dependent variable is  $RENO_t$ . Adjusted  $R^2 = 0.73$ ;  $F$ -Statistic = 61.97; Breusch-Godfrey serial correlation test (testing for up to 4<sup>th</sup> order autocorrelation) = 6.42; (critical value at 5%:  $\chi^2(4) = 9.49$ ); White test for heteroskedasticity = 0.68; (critical value at 5%:  $\chi^2(5) = 11.10$ ). The sample period is 1985:1–1996:2 and the number of observations is 46.

dependent variable was taken into account by explicitly introducing a time-trend variable in the estimation of Equation (7) but it was insignificant. Moreover, short-term influences of time reflecting seasonal variations were examined by using dummy variables, but the results did not indicate seasonal effects on the cyclical behavior of new retail orders. Finally, Wald tests for omitted variables showed that no significant information is lost in explaining the cyclical pattern of new retail orders by excluding the omitted lags of  $\Delta RENT$  and  $\Delta EX$  and changes in retail yields ( $\Delta RY$ ) from Equation (7).<sup>3</sup> Similar tests for the exclusion of the variables appearing in the Koyck model of new retail orders (Equation 5) and changes in retail sales both lagged one quarter, did not indicate that significant information was lost by not using these variables as regressors. Therefore, there is strong evidence that the model given as Equation (7) is superior in modeling new retail orders and the retail building cycle in Great Britain than the Koyck model of new orders (Equation (5)).

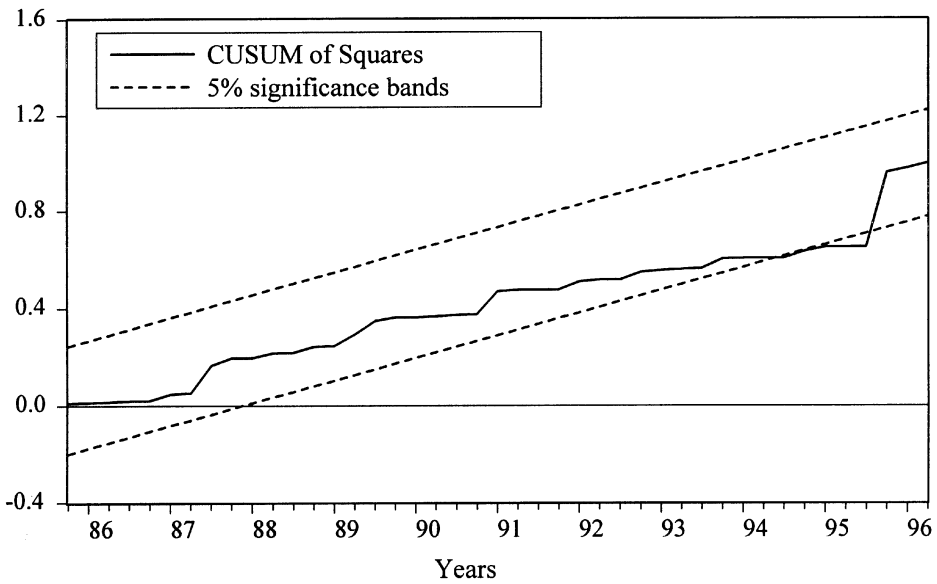
The actual and fitted values are presented in Exhibit 6, along with the variation in the residuals. The fitted values are the predicted values from Equation (7) computed by applying the regression coefficients to the independent variables ( $\Delta RENT_{t-3}$  and  $\Delta EX_{t-4}$ ). The fitted values seem to capture these general trends, but not the peaks and troughs. This is not surprising given the considerable short-term oscillations in the new retail orders series. The behavior of the residuals is overall considered satisfactory as the relevant tests indicated. However, an upward trend can be observed since 1995:1 and a spike in 1995:4. The residuals have exhibited such trends in past periods (*e.g.*, 1986:1 to 1987:3), but they reverted back to the zero mean line.

The upward trend in the residual values since mid-1995 requires an examination of whether Equation (7) has undergone a structural change before and after mid-1995, that is, whether the parameters of this equation have changed. The cumulative sum of squares (CUSUMSQ) test, which looks at the cumulative sum of the residuals, was used to find parameter instability. The plot of the cumulative sum of the squared residuals against time is given in Exhibit 7. Parameter instability occurs if the

**Exhibit 6**  
**The Actual and Predicted Retail Building Cycle**



**Exhibit 7**  
**CUSUM of Squares Test for Equation (7)**



cumulative sum moves outside the area between the two critical lines. In Exhibit 7, there is a movement outside the lower critical line but the cumulative sum reverts back to the area between the critical lines. This could be interpreted as an indication of parameter instability. Additional evidence on a possible structural change in the second half of 1995 is produced by the popularly used Chow test (Chow, 1960). The data set was split into two subperiods: 1985:1–1995:3 and 1995:4–1996:2. This test showed that a structural break occurred in 1995:3. The computed  $F$ -value was 6.74 and the Chow test is significant at the 1% level. Thus, there is evidence that the parameters of Equation (7) have changed between the periods 1985:1–1995:3 and 1995:4–1996:2.

This break could be attributable to a changing economic relationship between new retail orders ( $RENO$ ) and changes in retail rents ( $\Delta RENT$ ) or changes in expenditure ( $\Delta EX$ ) in the two time periods (before and after 1995:3). Consequently, a further test is undertaken to examine the stability of the coefficients on each of the regressors ( $\Delta RENT_{t-3}$  and  $\Delta EX_{t-4}$ ) of Equation (7) in order to determine whether the estimates of the coefficients differ significantly before and after 1995:3. For this purpose,  $RENO$  is regressed separately on each of the independent variables of Equation (7) (that is  $RENO$  on  $\Delta RENT_{t-3}$  and  $RENO$  on  $\Delta EX_{t-4}$ ) over the sample period 1985:1 to 1995:3. Then the remaining three observations (1995:4 to 1996:2) are added and the two regressions are run again over the whole sample period. An  $F$ -ratio is formed and the computed value is compared with the theoretical value of  $F$  adjusted for the appropriate degrees of freedom, to examine whether the coefficients are stable when the sample increases. The results for the coefficient estimates of  $\Delta RENT_{t-3}$  and  $\Delta EX_{t-4}$  are given in Exhibit 8. As regards the coefficient of  $\Delta EX_{t-4}$ , the observed  $F$ -value at the significance level of 5% is 1.71, which is smaller than the critical value of  $F$  at the same level of significance ( $(F(3,41) = 2.82)$ ). Therefore, the null hypothesis that the coefficient of  $\Delta EX_{t-4}$  is stable and not sensitive to sample size is not rejected. On the other hand, the test shows that the coefficients of  $\Delta RENT_{t-3}$  are unstable (different) in the two samples (before and after the three observations were added).<sup>4</sup>

It can, therefore, be argued that a breakdown in the estimated relation of new retail orders and changes in consumer expenditure seems to have occurred in 1995:3. However, what appears to be a structural break in the estimates could be a short-term disturbance in the value of new orders which is not captured by Equation (7). The volume of new retail orders increased by 58.1% between 1995:3 and 1995:4 and

**Exhibit 8**  
**Structural Stability Tests for  $\Delta RENT$**   
**and  $\Delta EX$**

Coefficient	$F$ -test	Critical Value $F_{0.05}(3,41)$
$\Delta RENT_{t-3}$	6.81	2.82
$\Delta EX_{t-4}$	1.71	2.82



subsequently fell by 42.5% between 1995:4 and 1996:1. These percentage changes are the largest in the whole sample period and since Equation (7) does not replicate them it means that other factors, in addition to changes in rents and expenditure, played a significant role in determining the values of new orders in the second half of 1995. It is important to point out that these influences were relevant in the second half of 1995 and were not persistent throughout the sample period as the diagnostics tests indicated. It is also worth noting that in conducting the Chow test, the significant increase in the volume of retail development starts in 1995:4 affects the size of the *F*-ratio used to test for parameter instability. Taking also into account that the CUSUMSQ test line reverts back within the critical bands following the break in 1995:3 it can be inferred that the results may be indicative of parameter instability in Equation (7) but this is not established. Moreover, since there are only three observations after the indicated structural break, a more rigorous examination can only be made when additional observations become available.

## Conclusion

The theoretical model of the retail building cycle assumes that the cyclical variation in new retail development in Great Britain, proxied by the volume of new development starts (new retail orders), is caused by changing trends in real consumer spending, real retail rents and real retail property yields. Granger causality tests to examine the assumed causal structure of the model indicated that there is no relationship between retail yields and the cycle of retail development. This suggests that investment market influences are either not captured by retail yields or cannot be examined at the aggregate level of analysis. A disaggregated analysis might be more pertinent to the examination of investment influences on retail property production, which is possible if investment influences are confined to particular geographical markets and types of retail property. Consequently, the retail building cycle was modeled on changes in real retail rents and real consumer expenditure.

The distribution of the expected lagged effects of these variables on retail development is captured by an Almon polynomial lag procedure which implies that retail property developers do not react immediately to changes in expenditure and rents. The preferred dynamic specification to explain the development cycle incorporated changes in rents lagged three quarters and changes in expenditure lagged four quarters. This model has a good explanatory power and a number of tests demonstrate its statistical robustness over the sample period. However, in the second half of 1995 a possible structural break in the proposed model was evaluated; there appeared to be a change in the economic relationship between new retail orders and retail rents. It was concluded that the observed break could have merely been the result of the abnormal changes in the level of new retail orders in the period 1995:3 to 1996:1 as a number of stability tests pointed to time invariant regression coefficients.

The present study generates evidence on a number of issues that are relevant to modeling the retail building cycle and property analysts. The aggregate retail building cycle largely reflects the cyclical variation in the demand for retail space. For the most part, the quantity of demand for retail space increases and decreases with the

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rise and fall of general economic activity. The results provide support for this contention and suggest that consumer expenditure is a significant indicator of the effects of aggregate economic conditions on the demand for retail space and provide the stimulus for initiation of new projects at least in Great Britain. Changes in consumer expenditure in the last six to eighteen months (two to six quarters) have an effect on the variation in the current level of new orders. The peak effect is found after approximately four quarters. Changes in real retail rents also affect the retail building cycle in Great Britain. The significance of this variable suggests that new retail development will be stimulated when demand pressure on the existing supply of space is sufficient to raise retail rents. Retail rent changes in the last three to fifteen months convey important information about the prevailing and expected conditions in the retail market, which are expected to influence the directions of the development cycle. Changes in real retail rents three quarters in the past exert the most significant influence on the current level of new retail orders. Given that the regression model is well specified, its estimated coefficients can be used to predict the future trend of the retail development cycle from current and expected expenditure and retail rent levels.

On the evidence of this study, the dynamic adjustment in the production of retail space implied by the Koyck distributed lag model, which was the basis for similar analyses in the U.S., does not model retail building development very well in the U.K. This may reflect different market characteristics and property investment decision processes or even differences in the measurement of the data series used. Therefore, real estate analysts should consider alternative specifications and be conscious of unique market adjustment processes when they examine retail property cycles and forecast in different markets around the world. The predictability of the retail building cycles is an area for further research. The output of such research work is of interest to those involved in retail property development and investment. Future research should identify the key forces that are responsible for the cyclical pattern of retail development in international markets. In particular, attention should be devoted to the transmission mechanisms and characteristics of these markets in order to pinpoint the sources that can lead to an oversupply of retail space and a plunge in retail values.

## Appendix

*New Retail Orders (RENO)*: New retail orders relate to contracts for new retail construction work awarded to main contractors by clients in the private sector including extensions to existing contracts and speculative work undertaken on the initiative of the firm where no contract or order is awarded. Retail construction work includes shopping centers, out of town stores and small shop units. The value of this work (in sterling) is recorded in the period when foundation works are started on retail projects for eventual sale or lease. The data on the value of new retail orders placed with contractors do not include other works incidental to the retail development such as infrastructure works (*e.g.*, road improvements). These figures are collected by the Department of the Environment in current price terms and published in *Housing and Construction Statistics*. For the purpose of this study they are revalued at constant

prices to produce a series of the volume of new retail construction in Great Britain and indexed (1990:2 = 100). Starting date of series: 1985:1.

*Total Consumer Expenditure (EX)*: Total consumer expenditure in Great Britain covers all personal expenditure on goods (durable and nondurable) and services. The data are compiled by the Office for National Statistics and are available at constant prices. The main consumers' expenditure headings are: cars, motorcycles and other vehicles; furniture and floor coverings; other durable and household goods; food; drinks and tobacco; clothing and footwear; energy products and other goods. The methodology followed in the production of the consumer expenditure series is outlined in full in *Consumer Trends*, a quarterly publication of the Stationery Office. The Office for National Statistics combines information from several sources to compile the data and the series is not based on an extrapolation of sales taxes. A main element in the collection of the expenditure data is a nationwide survey of households in which informants are asked to keep diary records of daily expenditures for two weeks. The survey is continuous and is supported by interviews that are spread evenly over the year to ensure that seasonal expenditure changes are covered. The data are indexed (1990:2 = 100). Starting date of series: 1955:1.

*Retail Rents (RENT)*: The series for retail rents is provided by CB-Hillier Parker international property consultants. Rental values relate to fully leased properties in good locations let of full repairing and insuring lease. The rent survey includes shop units located in the prime 100 per cent trading area of the town (high street) and shopping center in 500 locations throughout the country. The survey also extends to cover units located in the adjacent trading areas to high streets and shopping centers. It also includes retail warehouses defined as buildings well situated with modern specifications from which a wide variety of bulk goods are sold (gross floorspace of about 10,000 sq ft or more with a car parking ratio of at least 5 spaces per 1000 sq ft gross). The retail rent data are available as an index (1977:2 = 100). In this study, they are adjusted for inflation using the GDP implicit price deflator and re-indexed (1990:2 = 100). Starting date of series: 1977:2.

*Retail Property Yields (RY)*: Retail property yields refer to average gross income yields (expressed in percentage terms) achieved at the time of lease. Retail property yield data are also provided by CB-Hillier Parker. The survey is based on the same sample as that for the compilation of rents. The nominal yield data are expressed in real terms by subtracting the year on inflation rate each quarter from the quarterly nominal yields. Starting date of series: 1977:2.

*Retail Sales (RS)*: The Retail Sales Index is compiled each month from a sample survey carried out by the Office for National Statistics on 5,000 businesses in Great Britain, including large retailers and a representative panel of smaller businesses. The survey provides estimates on the volume and value of retail sales for all groups of retailers and the following broad group headings: predominantly food stores, total predominantly non-food stores, nonspecialized stores, textile clothing and footwear

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stores, household good stores, other stores. The data are indexed (1990 quarter 2 = 100). Starting date of series: 1971:1.

## Endnotes

<sup>1</sup> This system for *RENO* and  $\Delta EX$  takes the following form:

$$RENO_t = \sum_{i=1}^m a_{1i} RENO_{t-i} + \sum_{i=1}^m \beta_{1i} \Delta EX_{t-i} + \varepsilon_t$$

$$\Delta EX_t = \sum_{i=1}^m a_{2i} \Delta EX_{t-i} + \sum_{i=1}^m \beta_{2i} RENO_{t-i} + u_t$$

The tests undertaken aimed to examine whether all the coefficients of the lagged values of  $\Delta EX$  in the first equation are zero and whether the coefficients of lagged *RENO* can be considered to be zero in the second equation. Similar regressions are run and null hypotheses tested for *RENO* and  $\Delta RENT$  and *RENO* and  $\Delta RY$ .

<sup>2</sup> The deletion tests comprised three criteria: (1) lagged values of the regressors that did not take the a priori expected sign were dropped; (2) following the suggestion by Haitovsky (1969), variables whose *t*-Statistic was less than 1 were not included in the model, since by dropping these variables the adjusted  $R^2$  increases; and (3) the critical *t*-value was used as an additional guide in determining whether a variable was a candidate for exclusion.

<sup>3</sup> These results are available from the author on request.

<sup>4</sup> The observed structural break could also be due to a difference in the values of the constant in the two periods. To test for this possibility the dummy variable approach put forward by Cujarati (1970a,b) was used. This method did not show that the constant is different in the two periods. On the other hand, it indicated that the source of this break is the differential slope coefficient on  $\Delta RENT$ .

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