Regime shifts in ex post UK commercial property risk premiums

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Regime shifts in ex post UK commercial property risk premiums

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Using a Markov Switching Model, the hypothesis that ex post commercial sector risk premiums have stable mean values within a time-varying framework is investigated. The probabilities of shifting expected values and the transitional probabilities of remaining in a high (low)-risk state at each point in time were estimated. Results suggest that industrial and retail sectors exhibit regime shifting behaviour although the probability of shifting between high- and low-risk states, while significant, was low compared to them remaining the same. Investigation of the transitional probabilities suggested the propensity to shift regimes differs between sectors, but is generally more prevalent in periods of relative uncertainty.

Keywords: UK commercial property; ex post risk premium; regime shift; Markov Switching Model

1. Introduction

The behaviour of risk premia ex post as measured by the return on property over a relatively safe rate is a key element in property pricing and investment as it influences not only the anticipated rate of return, relative to other investments, but also factors such as the holding period and the exit yield to the investor (Baum & Crosby, 2008). While the evidence appears to suggest that observed property risk premiums are time varying, there has been little investigation into the nature, and particularly, the stability characteristics of such property risk premiums. In general, the maintained hypothesis has been that risk premiums, while they are likely to vary over time, will do so around a stable mean value which itself is invariant to changing economic conditions. Hence the possibility of changing mean values has typically not been explicitly accounted for in the property empirical literature (Dunse, Jones, White, Trevillion, & Wang, 2007; Key, Zarkesh, MacGregor, & Nanthakumaran, 1994; Wilson & Okunev, 1999; Wilson, Okunev, & Zurbruegg, 2004). In statistical terms, property risk premia are assumed to follow a linear stationary process whereby symmetry and linearity restrictions are imposed on the data generating process. As pointed out by Bodman (1998), the issue of whether

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economic time series exhibits asymmetric and non-linear cyclicality is critical for generating accurate estimates in the estimation, testing and forecasting of the series under investigation. Given the importance of property as an asset class and the importance of the time-varying characteristics as analytical tools for both institutional and private investors, this paper seeks to empirically investigate the time-varying characteristics of observed risk premia in the UK commercial property sector and to estimate a model which explicitly allows for structural changes and transition periods between changing market regimes.

There have been numerous studies in the literature attempting to capture the level and changes in property risk premiums using various empirical methods ranging from rules of thumb, through surveys, to various types of econometric analysis, the results of which tend to differ according to market and time period analysed as well as method of analysis employed. For example, historically the risk premium for the UK property is often quoted as being 2–3% over conventional long-dated gilts. Dubben and Sayce (1991) refer to 2%, Hargitay and Yu (1993) 1.5–2.5% and Mackmin (1995) 1–2%. Baum and Crosby (2008) refer to risk premiums as high as 4% in 1999, falling to an average of 3% in 2003/2004. In the decision of the Lands Tribunal in Cadogan Estates vs. Sportelli, a risk premium of 4.5% was determined for the UK housing, which might suggest a higher figure for commercial property.

Using an ex ante framework, Chau (1997) studied the commercial real estate risk premium in Hong Kong pre and post announcement of the repossession of Hong Kong by China and found that the pre-announcement risk premium for offices, retail and industrials to be 3.1, 3.8 and 6.0%, respectively. The ‘shock’ of the announcement doubled the risk premium for offices and retail and increased the industrial premium to 9.1%. Based on a survey of the US real estate investors, Shilling (2003), reported an ex ante risk premium over the period 1988–2002 of 6–6.75%.

Blundell (2009) utilising UK IPD data undertook an analysis of past returns for property and bonds from 1981 to 2008 which demonstrated that the risk premium, as measured by the property initial yield (6.4%) plus net income growth (6.3%), less depreciation (2.3%) and less gilt yields (7.3%) produced an average figure of 3.1%. However, depending on the state of the market, the risk premium varied significantly around this long-term average. For example, in mid-2007 the premium was of the order of 1.6% reflecting an initial yield of 4.6%, expected income growth of less than 5%, gilts at 5.3% and 2.5% for depreciation. Blundell himself highlighted the limitations of such analyses as representing only a snapshot at a point in time hence not having the ability to demonstrate the impact of changing market conditions. In order to address this static problem, the Blundell model considers trends in index-linked gilts, inflation and rental growth from 1982 to 2006 reporting in particular that the high level of investor optimism played a key role in determining yields in 2006. Similar results using the Schiller’s cyclically adjusted P/E method indicated a role for income expectations in determining investor excess returns. The author also suggests debt leverage, changes in lease structures and climate change are also likely to impact on property market risk premiums.

Tarbert and Marney (1999) using data covering 1978–1995 examined actual (ex post) values of commercial property risk premiums to determine whether the long-held belief of a 2% required risk premium was in fact achieved ex post. Their analysis showed that, over their sample period, the 2% traditional ex ante risk premium was not in fact delivered ex post and that the quarterly ex post risk premium
appeared to fluctuate around a value of zero, was mean reverting and that the volatility of risk premiums was variable. They also reported that *ex post* risk premiums were negative for short periods of time particularly during recessionary periods.

Other studies consider specifically the cyclical patterns in property markets. Wilson et al. (2004) for example, report evidence for securitised property risk premiums in a range of markets, suggesting that the cyclical pattern in these risk premiums provide signals that can be used as an investment tool by investors. Dunse et al. (2007) examine provincial office yields in the UK and review the changing city risk premiums relative to the City of London and relate this to weight of money invested in the provincial cities. As far back as 1982, Grebler and Burns, reported non-residential construction cycles in the USA of 7 years, while Wheaton (1987) reported US office building construction cycles of 20 years. Similarly, UK building cycles of between 4 and 20 years were identified by Barras and Ferguson (1985) using spectral techniques and Wilson and Okunev (1999), with real estate return data report evidence of cyclicality in the US, the UK and Australian markets. Key et al. (1994) also report cyclical patterns in all property returns with durations from 3 to 9 years.

The issue of cyclicality in actual property returns is critical for generating accurate estimates of risk premia. The extant literature suggests that required property risk premiums, being the market perception of the riskiness of real estate investments, vary considerably over time, driven by the interactions of a somewhat diverse mix of expectations of underlying economic factors and investor sentiment, which themselves change over time. What is unknown, however, is how sensitive expected values, as measured by mean values of observed risk premiums, are to changes in economic conditions. Essentially, over time economic events occur and during these episodes the dynamic behaviour of risk premiums might be expected to differ from what has gone before. It is through the investigation of this issue that this paper adds to the existing literature on property risk premiums.

First, it considers the UK commercial property sectors, namely, industrial, office and retail, all of which are likely to be relatively sensitive to underlying economic forces as reflected in the business cycle. The aggregate commercial property sector risk premium (all property) is also analysed to compare with the individual commercial sectors. Second, by utilising the Andrews-Quandt statistic, it conducts formal stability tests for the observed premiums within a time-varying framework and with unknown structural breaks. Third, by using a Markov regime-switching model (MSM) to analyse the dynamics of risk premiums *ex post*, it takes into account that many of the diverse factors which together drive premiums are (as discussed above) either unobservable and/or difficult to measure empirically. One can hypothesise therefore that the aggregate ‘forcing variable’ driving such changes in behaviour is likely to be unobservable. It is for example well known that the analysis of commercial property risk premium is complicated due, at least in part, to the heterogeneous nature of the asset class, the opaqueness of the market, the relatively low level of transactions and the localised nature of the markets. Fourth, the methodology utilised allows for the view that the types of shocks emanating from the broader economy during the cycle are likely to be unique to a particular time period – either positive or negative – and may therefore impact in a distinct way on individual sector risk premiums – all of which have implications for modelling, investment decisions and asset allocation.
In exploring these issues, Section 2 provides a theoretical and practical background by considering yield construction in property pricing, the data used and their sources. Section 3 outlines the main econometric methodology employed in the paper. Section 4 presents the results of the quantitative analysis while Section 5 offers conclusions.

2. Yield construction and data sources

2.1. Yield composition

The selection of the capitalisation rate is complicated, whether an initial yield or equivalent yield \( y_e \) is chosen, and requires the assessment of a range of systematic factors such as general economic conditions, finance rates and level of taxation, as well as specific risk perspectives, including for example, the sector, location, covenant strength of the tenant, unexpired lease length, letting risk, frequency of rent reviews and repairing terms.

Based on the work of Fisher (1930), the overall return on an investment is a reward for three factors: loss of liquidity, compensation for anticipated inflation and a risk premium \( p \). In order to simplify this equation, a risk free rate \( r_{fr} \) is used as a proxy for the loss of liquidity and anticipated inflation risk, with the redemption yield on conventional gilts usually adopted. These risk factors are offset to a greater or lesser extent by prospects for rental growth \( g \). Following on from the work of Baum (1988), an allowance for depreciation risk is added to compensate for both physical deterioration and obsolescence.

As the aim of this paper is to analyse dynamic properties of risk premia contained within the yield, the equation can thus be arranged:

\[
p = y_e - r_{fr} + g - d.
\]  

\((1)\)

In the empirical modelling reported below, each (decimalised) time-series observation of \( p \) was transformed into natural logarithms. Example for the office risk premium:

\[
\ln (1 + \text{officerrpg})
\]

where \( \ln \) denotes the natural logarithm.\(^3\)

2.2. Analysis period and data sources

The period analysed in this paper covers 1987Q1 through 2009Q4 with property yield data sourced from the Investment Property Databank (IPD) (www.IPD.com). The period of analysis was chosen for practical reasons: the IPD monthly index started in 1987 and the sample period captures periods of both rapidly rising and rapidly falling markets. As major economic indicators on the state of the economy are typically reported on a quarterly basis, it is reasonable to assume that quarterly intervals adequately capture the dynamic nature of the risk premium more effectively than those sampled on an annual or monthly basis.\(^4\) The data on the UK property equivalent yields covers three sectors, industrial, office and retail, plus ‘All Property’. The IPD UK quarterly equivalent yield data series did not start until the late 1990s, so the IPD UK monthly series\(^5\) was adopted which started in 1987. The
quarterly data points were taken at the end of March, June, September and December of each year.

2.3. **Risk free rate and the treatment of inflation**

For an asset to be considered risk free, the actual return must always be equal to the expected return and thus there can be no default risk and no reinvestment risk (Damodaran, 2002). With regard to the former, it is generally accepted that there is a minimal risk of default on all issuances of the UK government bonds (gilts). However, there are a range of different maturities of gilts with significant differences in their gross redemption yields, stability, yield distribution and volatility, raising the question of which maturity of gilt is most appropriate to use in the analysis of the risk premium over our study period of 1987–2009. Over the analysis period gilt yields were affected by a range of macro-level factors including the level of interest rates, level and slope of the yield curve, the rate of inflation, exchange rate policy, market volatility and the level of government borrowing. To eliminate reinvestment risk, all reinvestment rates would require to be known in advance. This would seem to encourage the matching of investment horizons as say, for example, the task was to estimate the expected risk free rate of return on an asset over a 10-year period, the use of say 3-month Treasury bills would not be suitable, as there is no way of knowing in advance what the Treasury Bill rate will be in three months time, thus exposing the investment to risk. The coupons on all bonds require to be reinvested at rates that cannot be predicted in advance and therefore all coupon paying bonds suffer from this weakness. In such a scenario, Damodaran (2002) recommends that only the expected return on government issued zero coupon bond would fully eliminate the reinvestment risk.

The literature is not definitive on what maturity of risk free rate to use, although most do refer to government bond rates. Baum and Crosby (2008) suggest that the life of the bond and the holding period should match. Sayce, Smith, Cooper, and Venmore-Rowland (2006) suggest the range of possible gilt yields options, and include in their rationale the long-term nature of property investments, holding period and lease length. Geltner, Miller, Clayton, and Eichholtz (2007) suggest that as real estate investment is long term and that as the typical Discounted Cash Flow analysis is carried out over a 10-year horizon, 10-year T-Bonds should be used, adjusted for the yield curve effect, while Diermier, Ibbotson, and Siegel (1984) employ a risk free rate based on the average of London Interbank Offered Rate.

A survey by Hutchison, Fraser, Adair, and Srivatsa (2011) of the UK property investment fund managers and their advisors, found that the majority, but by no means all of the respondents, used the 10-year nominal gilt yield as their risk free rate of return. The rationale for this approach appears to be partly because the 10-year gilt rate is a recognised, widely available and easily understood benchmark, and partly because it matches the expected length of their holding period of between 5 and 7 years on average. However, the results were not entirely clear cut as for example, some investors used the 10-year gilt rate even if their expected holding period was 30 years or more. Furthermore, concern was expressed whether it was appropriate to use the current rate on gilts, particularly when rates in 2009/2010 had fallen to such a low level and it was suggested that perhaps an average rate either based on gilt yields over the last 10 years, or a forecast rate, might be more relevant. In addition to the survey of the UK property investment community,
the authors analysed the stability of 2, 5, 10, 20 and 30 year gilts over the period 1980–2010 and found that 10-year index-linked gilts have been the most stable.

Inflation is another factor which needs to be accounted for in choosing a risk free rate. Inflation risk is variable across the term structure of the yield curve due to the year-on-year forward-looking inflation expectations in the market. The yield on conventional gilt reflects compensation for the time value of money and expected inflation, plus a risk premium to compensate for unexpected inflation. The level of risk premium for unexpected inflation will fluctuate depending on factors such as stability of the market. Gilts must deliver a competitive return and market prices adjust to changes in interest rates and inflation. The use of conventional gilts thus introduces further issues which require the estimation of both types of inflation and thus the use of real risk free rates would simplify matters. Given all the issues raised above, the analysis in this paper will use the yield on the UK government 10-year index-linked zero coupon bonds as the proxy for the real risk free rate. The index-linked data were obtained from the Bank of England (www.bankofengland.co.uk).

2.4. Depreciation

Depreciation is defined by Law (2004), as quoted in Investment Property Forum (IPF) (2005), as ‘the rate of decline in rental/capital value of an asset over time relative to the asset valued as new with contemporary specification’ (p. 7). The depreciation rates adopted were taken directly from the results of the IPF (2005) study (see Table 1). Based on the analysis of 659 properties over a 19 year holding period from 1984 to 2003, the annual rental depreciation rates were as follows.

While it is acknowledged that the analysis period in the IPD depreciation study is different from the analysis period in this paper, there is overlap in 17 out of the 23 years. The authors of the IPF research flagged that the results needed to be treated with caution due to worries over survivor bias in the sample, limitations in the recording of capital expenditure and that the data was extracted from the IPD database, which reflects the class of property preferred by institutional investors. It is also acknowledged that the rate of rental depreciation maybe time varying, affected by the property cycle, the level of new building, the sector, specific location and the life cycle of the building, but the IPF study did not provide this level of analysis. It is acknowledged that a varying depreciation rate could have been used but the numbers are small and the authors believe this to be a reasonable proxy for long-term real depreciation.

2.5. Growth rates

The pricing model requires an adjustment for growth and quarterly rental growth was constructed from annual IPD rental growth data which is based on estimates

Table 1. Rental depreciation rates 1984–2003.

<table>
<thead>
<tr>
<th>Sector</th>
<th>Rate of rental depreciation % per year</th>
</tr>
</thead>
<tbody>
<tr>
<td>All property</td>
<td>1.0</td>
</tr>
<tr>
<td>Office</td>
<td>1.0</td>
</tr>
<tr>
<td>Industrial</td>
<td>0.6</td>
</tr>
<tr>
<td>Standard shop</td>
<td>0.1</td>
</tr>
</tbody>
</table>

of rental value. The growth rates were deflated using the corresponding UK inflation data and interpolated into quarterly observations. There are of course many methods available to interpolate low frequency data series into a high frequency series. However, because of the lack of information provided in the low frequency series (and therefore the inability to recover actual values) whatever method used can only be suggestive of higher frequency values. In this study, the method used was an Autoregressive (AR) matching sum process, whereby an AR model was utilised in the interpolation process. To distribute the higher frequency values across the quarterly periods, a dynamic programming algorithm was used thus providing non-constant high frequency values. Essentially, the distribution of the series is computed by changing the frequency to a higher one (quarterly) while maintaining the sum of each quarter to be the same as that in each original (annual) period. In order to ensure the construction of risk premiums were not sensitive to interpolation methods used and followed a similar time path to the original annual rental growth data, we compared a number of different interpolation methods. The results indicated that they provided very similar high frequency observations and that non-constant type matching sum methods were not significantly different. The result of this exercise is reported in the Appendix to this paper. While quarterly observations were used in the analysis, in order to get an overview of the rental growth estimated, Table 2 shows the average per annum real growth rate over the period 1987Q1–2009Q4.

3. Econometric method

3.1. Stability tests

While standard theory dictates that real estate value, as described by Equation (1) above, is drawn from a single distribution with constant moments, there is a debate in the literature which questions the validity of this assumption, arguing that value is driven by a mixture of underlying processes which characteristically result in structural instability and often bi-modal distributions reflecting differing regimes. Two (not unrelated) reasons can be offered as an explanation of these observations with respect to the UK commercial property market. First, as commercial rental income is typically secured by legal covenant (e.g. the lessor retaining the physical asset if faced with tenant default), the risk premium has as a component an (unobservable) embedded option on economic conditions (see, e.g. Smith & Brooks, 1999). Second, the observed tendency of commercial property value to be highly cyclical, results in it experiencing periods of, often asymmetric, booms and subsequent slumps in line with changing expectations on economic conditions (see, e.g. Lizieri & Satchell, 1997). Such features, it is hypothesised, lead to ‘regime-led’


<table>
<thead>
<tr>
<th>Sector</th>
<th>Average real rental growth % per year</th>
<th>SD</th>
</tr>
</thead>
<tbody>
<tr>
<td>All property</td>
<td>2.81</td>
<td>8.09</td>
</tr>
<tr>
<td>Industrial</td>
<td>2.59</td>
<td>7.22</td>
</tr>
<tr>
<td>Office</td>
<td>1.87</td>
<td>11.61</td>
</tr>
<tr>
<td>Retail</td>
<td>3.85</td>
<td>5.61</td>
</tr>
</tbody>
</table>

Source: IPD & ONS.
behaviour in risk premiums according to whether economic conditions are, or are expected to, improve or deteriorate.

In order to establish the stability of sector risk premiums, we compute stability statistics according to the Quandt-Andrews breakpoint test for unknown structural breakpoints in a single series equation. Using standard information criterion statistics, the appropriate equation was specified as an AR model of order two (AR(2)) which ensured white-noise residuals (see, e.g. Andrews & Ploberger, 1994).

The idea underpinning such tests is that a single Chow breakpoint test is conducted at every observation between two dates, $\tau_1$ and $\tau_2$ and summary statistics computed from these for tests against the null hypothesis of no breakpoints. The relevant statistics are computed from each individual Chow breakpoint test, namely the Likelihood Ratio (LR) $F$-statistic and is a comparison of the restricted and unrestricted sum of squares. The summarised statistics are the Max$F$, which is the maximum of the individual Chow $F$-statistics, the Exp$F$, which is the logarithm of the exponential $F$-statistic and the Ave statistic which is the simple average of the $F$-statistics. As the distribution of these statistics becomes degenerate as $\tau_1$ approaches the start of the sample or $\tau_2$ approaches the end of the sample, it is suggested that the ends of the sample are not included in the testing procedure. This ‘trimming’ of data is suggested to be between 5 and 15% at each end of the sample. The results reported in this work computed the statistics using 5, 10 and 15% symmetric trimming.

3.2. The Markov Regime Switching Model (MSM)

Given the above discussion on property risk premia stability, it can be argued to be prudent to allow for characteristics such as structural shifts in mean when they are present, in the form of ‘regime-switching’ models. This is particularly the case when the use of linear models would incur substantial efficiency losses by the use of sub-sample data. There are of course different methods of accommodating regime shifts into the analysis, such as Threshold Autoregressive Models (TAR) (see, e.g. Tong, 1990) and the Markov Switching Model (MSM) due to Hamilton (1989, 1990). In this work, we select the latter for the analysis, motivated by the fact that the TAR model requires the shifts in regime to be directly observable, while the MSM method does not, thus assuming there is no single and/or observable variable driving the state of the risk premium. In contrast, the MSM makes probabilistic inference about shifts through the observed behaviour of risk premiums themselves. Given above discussions on the many interrelated factors associated with changes in risk premia, a TAR model would be difficult to justify in this instance. Further, as we have no knowledge of which process is likely to describe the data, the intuition is that regime change itself should be viewed as a random variable. Hamilton (1990) postulated that such a process provided a method of describing the probability laws governing alternate shifts in regime. In what follows, a brief discussion of the MSM is provided.

MSM’s have become popular in recent years mainly as the result of Hamilton (1989, 1990) and have largely been used to analyse financial asset and macroeconomic time-series data such as equity ratios (Brooks & Persand, 2001) asset returns (Bergman & Hansson, 2005; Calvet & Fisher, 2008; Smith & Brooks, 1999) and output growth (Hamilton, 1990, 2010). In general, these studies have found evidence of significant regime switching behaviour in these series advocating that such
models are useful when it is thought a series will shift from one type of behaviour to another and back again, but where the driver of this process is unobservable and/or difficult to measure. Clearly the application of a regime-shifting approach to risk premium behaviour is particularly useful if such premia have been identified as having the potential for structural instability.

The MSM is well documented.\(^9\) In its simplest form, the MSM assumes that there exist \(n\) states of nature corresponding to \(n\) regimes, with the mean shifting regime from one state to another and back again according to some unobservable process with the movements in the unobservable or state process governed by a Markov process: hence the probability distribution of the state at any time, \(t\), depends only on the state at \(t-1\) and not of those at \(t-2, t-3\) .... It is therefore able to capture changes in mean values between states and forecast the probability that it will be in a particular regime in the next period as well as a set of transition probabilities providing the probabilities of being in a certain state at a point in time.

In the case of two regimes, the transition probabilities can be expressed as:\(^{10}\)

\[
\begin{align*}
\text{Prob}\left[z_t = 1|z_{t-1} = 1\right] &= p_{11}, \\
\text{Prob}\left[z_t = 2|z_{t-1} = 1\right] &= 1 - p_{11}, \\
\text{Prob}\left[z_t = 2|z_{t-1} = 2\right] &= p_{22}, \\
\text{Prob}\left[z_t = 1|z_{t-1} = 2\right] &= 1 - p_{22},
\end{align*}
\]  

\(z_t\) is the unobservable state variable driving regime shifts and evolving according to a first-order Markov process with \(p_{11}\) and \(p_{22}\) being respectively, the probabilities of being in regime 1 (2) given that the system was in regime 1 (2) at \(t-1\). \(1-p_{11}\) and \(1-p_{22}\) are therefore, the probabilities that the variable of interest (in our case the property risk premium) will change from state 1 (2) to state 2 (1) next period. The approach therefore estimates the likelihood of the property risk premium entering a new phase or turning point next period, given information on the current state. Given that the new phase may itself be non-permanent, it allows for the possibility that it will return to its previous state. The parameters in each state and the state at which the economy is at each point in time can be simultaneously estimated without any need to assume the break date or whether it is gradual or sharp and allows for multiple switches between the regimes. Given the Markov process, it is the current period’s probability and set of transition probabilities (Equations (2.1) through (2.4)) that will be able to forecast the probability that the variable will be in a given regime during the next period.

Following preliminary analyses, two regimes for risk premia are specified, expansionary (regime 1) and contractionary (regime 2) allowing for separate means in each regime. Regime 1 therefore is associated with weakening economic conditions while regime 2 is associated with strengthening economic conditions.\(^{11}\) Also included in the model is a simultaneously estimated second-order AR process (AR (2)) for each of the property sectors under consideration, thus providing evidence whether allowing for an AR switching process can capture any non-stationary
processes embedded in risk premiums estimated as linear functions. The model is specified as:

\[ P_t = \mu_{st} + \Phi_1 p_{t-1} + \Phi_2 p_{t-2} + \epsilon_t \]  

where \( p_t \) is the risk premium, \( s_t, (t=1, 2) \) are the two states following a standard 2-regime Markov process, and \( \epsilon_t \sim N(0, \sigma^2) \).

4. Empirical results

4.1. Preliminary statistics

Figure 1 displays the natural logarithm of all four ex post risk premium series. The graph indicates an expansion–contraction range for the sectors of 7 to −4%. All sectors fell dramatically between 1989 and 1993 and again at the end of the sample period. The office sector appears to be the most volatile of all sectors, with a further notable fall in the level of risk premium between 2001 and 2003. These results are consistent with the findings of Tarbert and Marney (1999) and are not surprising given that the UK was in recession from 1990 to 1993 and experienced a significant downturn in economic activity in 2000/2001. The recession in 1990 was particularly severe in the office sector, as it followed a boom in the market during the late 1980s. In the period 1987–1989, IPD reported total nominal annual property returns in the Office sector of +30.7, +31.1 and +16.5%, respectively. In the three subsequent years, nominal returns of −9.9, −10.8 and −7.2% were recorded as the sector suffered significant negative rental growth. This period saw Office equivalent

![Figure 1](image-url)
yields move out from 7.66% at the end of 1989, to 9.4% 3 years later, while the yield on nominal 10 year gilts fell from 10.25 to 6.09% over this period as investors sought refuge in the relative safety of government stock. Essentially, the equivalent yield did not adjust quickly enough to the reality of declining rental growth in the market place. To estimate the office sector ex ante risk premium from the realised performance of this sector relative to the safe rate would clearly have been nonsensical. It was in effect a very disappointing period for commercial property investors.

Table 3 reveals that the quarterly mean value of the risk premiums is highest for retail and industrial at 2.1% per quarter and lowest for office at 1.6% per quarter, with the all property figure at 1.8% per quarter. Consistent with Figure 1, office display’s the greatest ex post volatility with a standard deviation of 2.7% per quarter. For all series, the PP statistic cannot reject the null hypothesis that risk premiums are non-stationary. This suggests all series do not exhibit a tendency to revert to a single mean value.

However, it is by now well known that standard unit root tests have low power, particularly for near unit-root process. Perron (1990), for example, shows that unit root tests are biased toward the non-rejection of the null hypothesis (which is non-stationarity) in the presence of structural instability. Therefore, we can conjecture that a reason for the non-rejection of the null of non-stationarity reported in Table 3 is because of structural instability in mean values (see also Bergman & Hansson, 2005). This is a feature of the data we return to in the analysis below.

The Q-statistics for all series suggest the presence of autocorrelation in the series with the correlogram for all series indicating lags one and two as being statistically significant (not reported). While departures from normality as tested by the $J-B$ statistic are statistically significant for the industrial and retail sectors there is less evidence of this feature in the summary statistics for the other sectors. In an attempt to reveal differences in the sector distributional structures, Figure 2 compares the estimated distributions with a normal distribution with the same mean and variance.

Figure 2 indicates some evidence of skewness in the series where the estimated series display relatively longer right tails. For all property, industrial and retail, there is some evidence of right-hand shoulders to the distributions while for office, the evidence is less pronounced. Consistent with the summary statistics above, the industrial and retail sectors display the greatest deviations from normality with their


<table>
<thead>
<tr>
<th>Sector</th>
<th>Mean</th>
<th>SD</th>
<th>PP test</th>
<th>$Q$-stat (4)</th>
<th>$J-B$</th>
</tr>
</thead>
<tbody>
<tr>
<td>All property</td>
<td>.018</td>
<td>.018</td>
<td>−1.971</td>
<td>273.63 (.000)</td>
<td>3.680 (.159)</td>
</tr>
<tr>
<td>Industrial</td>
<td>.021</td>
<td>.017</td>
<td>−1.913</td>
<td>297.00 (.000)</td>
<td>30.427 (.000)</td>
</tr>
<tr>
<td>Office</td>
<td>.016</td>
<td>.027</td>
<td>−2.111</td>
<td>256.30 (.000)</td>
<td>1.010 (.603)</td>
</tr>
<tr>
<td>Retail</td>
<td>.021</td>
<td>.012</td>
<td>−1.839</td>
<td>271.79 (.000)</td>
<td>18.098 (.003)</td>
</tr>
</tbody>
</table>

Notes: SD denotes standard deviation and PP is the Phillips-Perron test for unit roots which is robust with respect to autocorrelation and heteroscedasticity in the equation disturbance process. The test equation included a constant and linear trend. Critical values for the PP test are: $-4.062$ (1%); $-3.460$ (5%) and $-3.156$ (10%). The sample period is 1987Q1–2009Q4. The $Q$-stat (4) is a test for autocorrelation in the series of up to 4 lags. $J-B$ is the Jarque Bera test for normality. Figures below the tests statistics are significance levels. Critical values for the PP test are: $-4.062$ (1%); $-3.460$ (5%) and $-3.156$ (10%). The sample period is 1987Q1–2009Q4.
shape suggesting two separate modes or regimes: the lower part of the distribution covering most of the observations and a higher part, capturing the highest values of the risk premiums.

As a precursor to formal stability tests and MSM estimation, we first consider the trending and cyclical behaviour of risk premia in our sample by decomposing each series into its trend and cyclical using a standard Hodrick–Prescott (H–P) filter which is a smoothing method to extract a smooth estimate of the long-term trend of a series revealing the cyclical component.14

Figure 3 displays the long-term trend of each series on the right-hand axis with the temporary cyclical component on the left hand axis. For industrial and retail, the patterns suggest that their temporary cyclical movements were more subdued from 1994 than they had been prior to this period when the UK economy was in deep recession. The office sector, with the exception of the mid-to-late 1990s, exhibits relatively high cyclical volatility over the sample. While in the later part of the sample, the graph indicates relatively constant, gently downward sloping trending patterns in risk premiums, this does not appear to be the case for the earlier part of the sample. Prior to 2000, risk premiums long-term trends also tended to vary over time with the time variation tending to be in tandem with those of the temporary or cyclical components. Overall, patterns in both temporary and long-term risk premium components suggest that risk premia are time varying and tend to be sensitive to ‘good’ and ‘bad’ economic conditions albeit such sensitivity appears to dif-
fer between sectors. Observed time variability however, does not necessarily imply statistically significant structural instability in risk premiums and in order to investigate this further, we proceed by reporting in Table 4 the Andrews–Quandt stability tests as discussed in Section 3 above.

It is clear from Table 4 and consistent with the preliminary results reported above, that the two candidates for structural instability in the UK property sectors over this period are the industrial and retail sectors: for these sectors the p-values imply that the null hypothesis of no breakpoints has been rejected. This is particular in the case for the industrial sector where all three statistics convincingly reject no structural changes in the model intercept and parameters.

Given the above results, the MSM model focuses on modelling risk premia for the two sectors where evidence of structural instability was identified, namely, the industrial and retail sectors. As discussed above, we assume two-multiple shift regimes each reflecting contraction and expansion of economic conditions respectively: state 1 = expansion of the risk premium and state 2 = contraction of the risk premium and compute the probabilities of remaining within/changing from, expansionary and contractionary regimes during the following period and their statistical significance. Table 5 reports the results of this exercise for these two risk premium series.

For the industrial and retail sectors, the two regime means, $\mu_1$ and $\mu_2$, are statistically significant and the chi-squared tests convincingly reject the null that risk premium mean values in the two states (premia expansionary and premia contractionary) are the same. As would be expected, the means in the high-risk state, $\mu_1$,
are higher than those in the low-risk state, \( \mu_2 \). The estimates of the AR coefficients with significantly positive AR(1) and significantly negative AR(2) parameters, sum to less than unity and, in contrast to the PP unit root tests reported above, suggest the risk premia series are stationary, therefore mean reverting, after allowing for structural shifts in mean values.

The estimated probability values, \( p_{11} \) and \( p_{22} \), (premia expansion and premia contraction respectively) suggest that for the industrial sector the probability of there being a shift (1\( - p_{11} \)) in the ‘high’ risk premium state is 10%, while during periods of risk premium contraction, this falls to .01%. This is also the case for the retail sector: the probability of changing states when in a high-risk state is 6.9% while it is 2% in premia contracting regimes, hence low-risk, states. Such evidence would support the view that there are asymmetries in the behaviour of risk premi- ums across regimes and across sectors.

The transitional probabilities of being in a high-risk premium state next period, at any point in time, are featured in Figure 4.

Consistent with the reported probabilities in Table 5, it is clear that the observed non-linearity associated with the industrial sector is due to one breakpoint peaking in 1988 with a duration of around 10 quarters with the subsequent period being

Table 4. Tests of structural breaks in UK property real risk premiums.

<table>
<thead>
<tr>
<th>Sector (%)</th>
<th>MaxF</th>
<th>ExpF</th>
<th>Ave</th>
</tr>
</thead>
<tbody>
<tr>
<td>All property</td>
<td>Data trimming</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5</td>
<td>5.651 ( p = .024 )</td>
<td>.945 ( .204 )</td>
<td>1.195 ( .271 )</td>
</tr>
<tr>
<td>10</td>
<td>4.561 ( p = .065 )</td>
<td>.532 ( .569 )</td>
<td>.830 ( .552 )</td>
</tr>
<tr>
<td>15</td>
<td>2.179 ( p = .587 )</td>
<td>.328 ( .824 )</td>
<td>.594 ( .766 )</td>
</tr>
<tr>
<td>Industrial</td>
<td>Data trimming</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5</td>
<td>9951 ( p = .000 )</td>
<td>2.745 ( p = .000 )</td>
<td>2.626 ( p = .008 )</td>
</tr>
<tr>
<td>10</td>
<td>9.274 ( p = .000 )</td>
<td>2.256 ( p = .006 )</td>
<td>2.237 ( p = .027 )</td>
</tr>
<tr>
<td>15</td>
<td>9.274 ( p = .000 )</td>
<td>1.495 ( p = .002 )</td>
<td>1.815 ( p = .080 )</td>
</tr>
<tr>
<td>Office</td>
<td>Data trimming</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5</td>
<td>4.403 ( p = .103 )</td>
<td>.074 ( .334 )</td>
<td>1.347 ( .193 )</td>
</tr>
<tr>
<td>10</td>
<td>4.403 ( p = .081 )</td>
<td>.730 ( .353 )</td>
<td>1.313 ( .217 )</td>
</tr>
<tr>
<td>15</td>
<td>4.403 ( p = .065 )</td>
<td>1.302 ( .228 )</td>
<td>1.952 ( .060 )</td>
</tr>
<tr>
<td>Retail</td>
<td>Data trimming</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5</td>
<td>7.350 ( p = .003 )</td>
<td>1.540 ( p = .039 )</td>
<td>2.432 ( p = .013 )</td>
</tr>
<tr>
<td>10</td>
<td>5.292 ( p = .029 )</td>
<td>1.130 ( .116 )</td>
<td>2.068 ( .040 )</td>
</tr>
<tr>
<td>15</td>
<td>2.984 ( p = .298 )</td>
<td>1.036 ( p = .144 )</td>
<td>1.952 ( p = .060 )</td>
</tr>
</tbody>
</table>

Notes: The model used in all tests is an AR(2) model. The tests are performed to test for shifts in the intercept and parameters of the model. The MaxF provides the maximum of the individual Chow-F statistics, ExpF is the Andrews-Ploberger Exponential and Ave is the simple average of the F-statistics. Data trimming refers to the recommended asymmetric shaving of observations at each end of the sample period. The number of breaks compared was 81, 73 and 63 for trimming values of 5, 10 and 15%, respectively. The figures in parenthesis under the statistics are \( p \)-values and are the Hansen (1997) approximate asymptotic \( p \)-values.
dominated by a low-risk regime with no expected switching. The retail sector however appears to have experienced three periods over the sample when regime changes were expected: two between 1987 and 1993 and the most recent and largest in 2009 Q2 at .98. This abrupt change in expectations occurs after a long period of slowly declining probabilities of experiencing high-risk regimes from a previous peak in the (recessionary) early 1990s. The number of probability spikes (signalling the likelihood of the risk premium rising) in the retail sector indicates relatively more uncertainly in this market – this of course is not surprising given the reliance of this sector on the sentiment of consumers and how this is reflected in actual consumer expenditure and therefore profits. For the industrial sector, and consistent with the estimated probability values from Table 5 above, the low-risk regime dominated the whole period from 1990 while for the retail sector the duration of perceived low risk was around 14 years.

A logical progression of the above analysis would be to formally compare the AR(2) linear model against the MSM non-linear specification by using standard LR tests. However, it is well known that the asymptotic distributions of such tests are affected due to the transitional probabilities not being identified under the null and the scores related to the alternative being zero under certain parameter values (Breunig & Pagan, 2004; Garcia, 1998; Hamilton, 1990). Unlike linear models however, the MSM model does predict that the non-linear transitional probabilities should have a positive and significant coefficient when added to an AR(2) representation (Hamilton, 1990). Once again however as the transitional probabilities are a generated series, it is not clear what distribution is appropriate for commenting on the significance of these coefficients. Given this constraint, in this instance, the appropriate regression equation is:

![Industrial and retail sector transitional probabilities](image-url)
Table 5. Estimates of a 2-state Markov switching AR(2) model for industrial and retail risk premiums.

<table>
<thead>
<tr>
<th>Sector</th>
<th>$\mu_1$ (expansion)</th>
<th>$\mu_2$ (contraction)</th>
<th>$\phi_1$</th>
<th>$\phi_2$</th>
<th>$\sigma^2$</th>
<th>$p_{11}$</th>
<th>$p_{22}$</th>
<th>$\chi^2_{(1)}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Industrial</td>
<td>3.590 (.169)</td>
<td>1.810 (.406)</td>
<td>1.855 (.005)</td>
<td>-.885 (.005)</td>
<td>.130 (.015)</td>
<td>.900 (.000)</td>
<td>.999 (.000)</td>
<td>19.066 (.000)</td>
</tr>
<tr>
<td>Retail</td>
<td>1.925 (.248)</td>
<td>1.644 (.257)</td>
<td>1.865 (.051)</td>
<td>-.897 (.048)</td>
<td>.080 (.006)</td>
<td>.931 (.102)</td>
<td>.980 (.019)</td>
<td>47.633 (.000)</td>
</tr>
</tbody>
</table>

Notes: $\mu_1$ (expansion) and $\mu_2$ (contraction) denote the model intercept, which is allowed to vary between the two states. $\phi_1$ and $\phi_2$ are the autoregressive (AR(2)) parameters and $\sigma^2$ is the variance. $p_{11}$ and $p_{22}$ are the probabilities of being in regime 1 (2) given that the system was in regime 1 (2) at $t-1$. $\chi^2_{(1)}$ is a chi-squared statistic with 1 degrees of freedom under the null hypothesis that $\mu_1 - \mu_2 = 0$. Figures in parenthesis under the estimated parameters are standard errors and under the chi-squared statistic are significance levels. Maximum Likelihood estimation was by Broyden Fletcher Goldfarb Shanno method.
where $p_t$ is the risk premium and $\text{prob}_t$ is the transitional probability of being in a high-risk premium at a point in time thus summarising the non-linear information from the previous period regarding the unobservable state variable. The $\gamma$ coefficient should therefore be positive and significantly different from zero. Table 6 reports the coefficient estimates for both the industrial and retail sectors.

As indicated above, it is difficult to comment definitively on the significance of the $\gamma$ coefficient estimates. However, using standard $t$-tests, and inconsistent with the predictions of standard linear models, both coefficients are, significantly positive indicating that (non-linear) probabilities, hence expectations, regarding shifting risk premium regimes have a significant role to play in the sector risk premium data generating processes.

5. Conclusion

The focus of this paper is on the dynamics of ex post UK commercial property sector risk premiums over the 23-year period from 1987 through 2009. In particular, it empirically investigates the maintained hypothesis that observed sector risk premiums have stable mean values within a time-varying framework. Using a MSM, the probabilities of shifting expected values and the transitional probabilities of remaining in a high (low)-risk state at each point in time were estimated. Evidence was provided suggesting that the industrial and retail sectors did exhibit regime shifting behaviour. However, the probability of changing mean values, therefore shifting between high- and low-risk states, while significant, was very low compared to them remaining the same. Investigation of the transitional probabilities also showed that the regime shifting was mainly concentrated in the early part of the sample although for retail, it was also observed at the end of the sample. Some evidence was also provided supporting the view that expectations regarding shifting regimes can have a significant role to play in the sector risk premium ex post data generating process.

The analysis therefore challenges the maintained hypothesis that ex post risk premiums, while likely to vary over time, do so around a stable mean value which itself is invariant to expectations regarding economic conditions. The findings support those who challenge the validity of the assumption that real estate value is drawn from a single distribution with constant moments, suggesting that value is driven by a number of underlying processes which characteristically produce structural instability and often bi-modal distributions reflecting the differing regimes.

In order to provide a greater understanding why there are differences between the sectors and why there is a relatively higher probability of shifts in the 'high' risk premium state for the industrial and retail sectors, further research will need to

<table>
<thead>
<tr>
<th>Sector</th>
<th>$\mu$</th>
<th>$\Phi_1$</th>
<th>$\Phi_2$</th>
<th>$\gamma$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Industrial</td>
<td>.001 (.000)</td>
<td>1.735 (.051)</td>
<td>−.787 (.049)</td>
<td>.111 (.025)</td>
</tr>
<tr>
<td>Retail</td>
<td>.000 (.000)</td>
<td>1.710 (.047)</td>
<td>−.737 (.047)</td>
<td>005 (.001)</td>
</tr>
</tbody>
</table>

Notes: Figures in parenthesis under coefficient estimates are standard errors. The sample period is 1897Q1 through 2009Q1.
be undertaken. The evidence from the current analysis suggests that investors may be adopting a different investment strategy between the sectors, return composition, holding periods or a combination of all three and indeed other possible factors. Investment in offices and in particular London offices for example, tends to be a yield ‘play’ with investors surfing the yield curve, holding the investment for relatively short periods of around 5 years (Gardner & Matysiak, 2005; MacGowan & Orr, 2008). While prudent timing of entry and exit from an investment is a relevant strategy across all sectors, prospects for income return is possibly the more dominant factor in the other sectors. Industrial and retail (shopping centres in particular) are sectors where a longer holding period is evident, with income the key determinant of return, which, in turn, is closely linked to changing economic conditions.

In addition, we have already identified that the findings support the view of real estate value being drawn from multiple distributions with variable moments. The probability of changing mean values across sectors may be influenced by expectations of return and risk within each sector and across the economic and property cycles. Again further research is required before we can draw definitive conclusions.

Overall, the empirical results reported above suggest that while shifting regimes do not appear to be so relevant in the commercial property sector as in equity, bond and output markets, it does exist in some sectors, particularly in periods of uncertainty about the wider economy. The evidence that commercial sector premiums display different data generating behaviour addresses, at least in part, limitations of earlier static analyses which do not take into account the impact on mean values of expectations on changing market conditions. The findings suggest that the simple averaging of individual sector risk premiums ex post may mask what can be significant structural changes, which in turn are likely to have important implications for diversification strategies and perceptions of required risk premia. Further, rules of thumb and traditional factor models which do not allow for the possibility of dramatic, albeit occasional, abrupt shifts in trend run the risk that reported results will be inaccurate and that trading rules and forecasts constructed from those results will not be well founded.

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Notes

1. Notably, in 2009, global invested stock stood at almost US$ 11 trn, with the UK property market accounting for US$ 562 bn (DTZ, 2010). Typically, a property investor would translate the current and expected risk premium into a target or hurdle rate for consideration in the investment decision process and, as inherent risk varies across asset classes due to asset-specific idiosyncrasies and uncertainty in their financial returns, an investor would also be interested in the relevant size and movements over time in risk premiums between asset classes.


3. It is well known that it is easier to derive the time-series properties of additive processes (continuously compounding by taking natural logarithms).

4. Annual data for example, would provide too few observations for a robust analysis using time series techniques, while the use of monthly data, which typically exhibits long-lived autocorrelations, would entail considerable adjustments for this in the time series modelling procedure.

5. The IPD UK monthly series is a subset of the main index and as at September 2010 contained 3,631 properties worth £32.9 billion of which £15.1 billion were in the retail sector, £10.2 billion in offices, £5.9 billion in industrial and £1.5 billion classified as ‘other’ which includes leisure properties. There were no residential properties. The ‘All Property’ returns series is a summation of the returns from the four sectors. For comparison, as at December 2009 the UK Annual Property Index contained 10,986 properties and was valued at £117 billion.

6. It is acknowledged that as indexing is done ex post, the ex ante risk free rate still has inflation risk.

7. Index-linked gilts differ from conventional gilts in that the semi-annual coupon payments and the principal are adjusted in line with the UK Retail Prices Index. This means that both the coupons and the principal paid on redemption of these gilts are adjusted to take account of accrued inflation since the gilt was first issued.

8. Rental value growth is calculated as the change in market rental value expressed as a percentage of the rental value at the start of the period.


10. In the case of regimes > 2, Equation (4) could be expressed succinctly in matrix form.

11. While the imposition of two regimes associated with structural changes is intuitively appealing, higher order regimes were also investigated. The coefficients on the mean parameter of regimes > 2 had large standard errors and could not be rejected as being equal to that of the first regime.

12. As indicated above an AR(2) process for all risk premiums in the sample was found to be optimal according to standard information criteria.

13. It is also well known that the realised equity premium has varied over time and, as pointed out by Mehra and Prescott (2008), the variation depends on the time horizon over which it is measured. Several authors report periods when the ex post equity risk premium has been negative. For example: 1991–1992 and 2000–2003 (Mehra & Prescott, 2008); and 34 individual years between 1931 and 2001, including 1992–1994 and again 2000–2001 (Dimson, Marsh, & Staunton, 2003).

14. While there are alternatives to the H-P filtering algorithm, the aim here is simply to use a consistent de-trending method across the sector risk premiums in order to compare their trending and cyclical properties. None of the results that follow are based on filtered data.

15. Research by Hutchison, Adair, & Findlay (2011) on covenant strength found that some sectors are likely to have a higher probability of insolvency and delinquency in a recessionary market and that a higher risk premium should be applied to those sectors that are more volatile.

References


**Appendix**

There are many methods available to interpolate low frequency data into high frequency observations. However, because of the lack of information provided in the low frequency series (and therefore the inability to recover actual values) any method used can only be suggestive of higher frequency values. Figure A1 shows the quarterly property real rental growth series resulting from three different methods of the interpolation of the annual real rental growth rates to a quarterly frequency (per cent per quarter) while Figure A2 displays the original annual series (per cent per annum). The methods used were: constant match sum
Figure A1. Frequency conversion: a comparison of methods of interpolation.

Figure A2. Annual real rental growth rates (%).
which assigns a constant value so that the sum of the high frequency observations matches the low frequency observation; quadratic match sum fits a local quadratic polynomial for each observation in the low frequency data, and uses this to provide high frequency observations for the specific period; AR match sum, whereby an AR model is used in the interpolation process, which is solved using a dynamic programming algorithm, to distribute the high frequency values. As with the quadratic method, the interpolated values, while they sum to the periods low frequency value are not constant.

As indicated by the graphs there is little difference between the methods used in terms of values and the time paths, with the exception that the constant match sum method has, by construction, a stepped time path while the others are smoother, following more precisely the smooth profile of the original data. The annual mean and volatility values were the same as those for the annual data. Construction of risk premiums using either the AR matching procedure or the quadratic matching procedure were not sensitive to the type of interpolations method used while, as expected, the constant sum method resulted in a ‘stepped’ series – see Figure A3 which displays, as an indicative example, the all property risk premium calculated using the three types of interpolations. Further as the quadratic method is recommended to be more suited to a situation where only a few data points are being interpolated, the AR matching sum model was chosen as being appropriate for this study. Other methods were also investigated such as constant match average; cubic match last; linear match last; quadratic match average, all of which gave very similar time paths as above when converted into quarterly data.