ESTIMATING BANKS' EQUITY DURATION: A PANEL COINTEGRATION APPROACH

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Abstract

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1. Introduction

This paper examines the issue of bank equity duration. Duration is a very useful concept. Among others, it provides an indication of how fast an investment is recovered and also, how sensitive prices are to interest rates. The contribution of this paper lies in two areas – in its analysis of the sensitivity of banks equity prices to interest rates, and in demonstrating the use of a new approach to empirically estimate duration. The sensitivity of share prices to interest rate changes, particularly in relation to banks, is one that is of great importance to investors and policymakers. Such information provides indications for investors of the profitability of their share portfolios, and for policymakers such as the Central Bank, of the impact of monetary policy. It is not clear yet in the literature as to whether or not bank share prices are sensitive to interest rates. In this paper, we provide further evidence on this issue with reference to four countries – Australia, US, UK and Canada.

In terms of the methodology in estimating duration, at present, there is also no clear agreement. There is the classical or traditional approach based on the McCauley model as well as a number of other approaches such as those of Leibowitz (1986). These models, given their own strengths and weaknesses, provide conflicting estimates of duration. As a compromise, it has been suggested that duration should be estimated empirically. We therefore do this in this paper. We empirically calculate duration using unit root and cointegration based on panel data. It is well-known that the use of panel data results in more information being taken into account because the cross-sectional dimension is combined with time-series dimension. In addition, this approach allows for individual effects and comparison of duration for the bank sector of different countries within the same model.

2. Banks' Share Price Sensitivity to Interest Rates: Evidence

Studies have found that the stock market in general significantly reacts to interest rate changes (see, for instance, Fama and French, 1989 and Fama, 1990). However, the case of bank stocks is not that clear. Since banks are in the business of borrowing and lending money, they are expected to be more sensitive to interest rates than non-financial firms. There is evidence to support this claim. A number of studies have found that bank stock returns are more sensitive to interest rates than stock returns of non-financial firms (see, for instance, Lynge and Zumwalt, 1980; Flannery and James, 1984 and Booth and Officer, 1985). Studies on the short-run impact of interest rate on bank share prices also showed that bank shares are sensitive to interest rate changes, as demonstrated for instance by Kaen, et. al. (1997) in their study of the price effect of Bundesbank interest rate changes during the period 1985 to 1983 on German bank stocks. Sweeney and Warga (1986), Kane and Unal (1988) and Yourougou (1990) also found evidence of bank shares' sensitivity to interest rate changes with the latter two pointing out that this sensitivity is time varying. Further studies on the time varying nature of the sensitivity of bank stocks to interest rate changes have been undertaken based on the use of ARCH-based models. Results of these studies revealed that both the levels and volatility of interest rates significantly affect the first and second moments of bank shares. Examples of these studies are those of Song (1994), Flannery et. al. (1997) and Elyasiani and Mansur (1998).

On the other hand, since banks are able to hedge their exposure to interest rate changes, this may make them not sensitive to interest rates. Banks can hedge against interest rate risk through the matching of their assets and liabilities – hence, their interest rate sensitivity could be dependent on the extent of mismatch in maturity between their assets and liabilities. The nominal contracting hypothesis developed by French et. al. (1983) postulates a direct relationship between bank equity sensitivity to interest rates and the amount of net nominal assets (nominal assets minus liabilities) and the duration of those assets. There is a view, however, that while it is logical that maturity mismatching between assets and liabilities can make the banks more exposed to interest rate risk, this

may not necessary affect the bank's returns. This is because banks can earn positive spread between deposit rates and lending rates and this spread could compensate for the interest rate effects. In other words, if the increases in interest rates on short-term funds can be passed on in terms of increases in long-term lending rates, then the profitability of banks will be unaffected by increases in interest rates and therefore bank shares will not be sensitive to interest rate changes. This contention is supported by several studies such as that of Flannery (1983), which found no impact of interest rate changes on bank stocks in the long-run, and those by Flannery and James (1984), Aharony, et. al. (1986) and Akella and Greenbaum (1992).

3. Duration Measurement and Estimation: Methodology

The concept of duration was initially developed by Macaulay (1938) and Hicks (1939) as a measure of the average number of years to recover a loan's present value. However, Hicks (1939) also showed that it is also essentially an elasticity measure of the values of capital asset with respect to the interest rate (discounting factor). Duration, therefore, is a measure of the relationship between the price of an asset and interest rate. In the case of bonds, this is relatively straightforward. From the bond valuation formula, it is clear that the bond price is inversely related to the yield to maturity. But in the case of equities, this is not so since there are many more factors, aside from interest rate, that impact on the price of a stock. This is easily seen from the dividend discount model (DDM) where the price of a stock is a function of its cash flows and the discount factor. The cash flows themselves can be dependent on a number of factors in which interest rate could be one of them. The traditional measure of duration, based on the dividend discount model, is shown below (see Boquist, et. al 1975):

$$D = \frac{(1+k)}{(k-g)} \quad (1)$$

where k is the discount rate and g the growth rate in dividend.

Leibowitz (1986) proposed another way to calculate equity duration based on the combination of the following equations pertaining to equity returns and bond returns:

$$R_E - R_f = a_1 + \beta_E (R_B - R_f) + e_1 \quad (2)$$

$$R_B - R_f = a_2 + \beta_B (\delta) + e_2 \quad (3)$$

where:

R_E $\equiv$ returns on equity market,

R_B $\equiv$ returns on bond market,

R_f $\equiv$ risk - free rate,

δ $\equiv$ change in benchmark long - term yield,

β_E $\equiv$ sensitivity of equity market returns to bond market returns,

β_B $\equiv$ sensitivity of bond market returns to benchmark long - term yield.

The traditional measure of duration uses price while that of Leibowitz (1986) utilises return, so they produce different duration figures. The latter results in lower duration because reinvestment risk offsets price risk, as pointed out by Johnson (1989). DDM duration has been found to yield much higher value than that of the Leibowitz (1986) model (Leibowitz and Kogelman, 1993, p. 51); thus, there is what is called an equity duration paradox. Hurley and Johnson (1995) attempted to reconcile the difference between the two measures of duration by exploring the traditional or DDM model further in terms of allowing a non-linear path for dividends in a kind of Markov sense, thus adding the convexity property of duration into their model. Their models produced equity duration magnitudes that are in between those of the traditional and Leibowitz models. Leibowitz and Kogelman (1993) also made such an attempt but again their results differ substantially from the models mentioned earlier. Hence, there is still no agreement as to which method of calculating duration is best.

This means that there is still scope for contribution to the literature on bank sensitivity to interest rates and on the measurement of duration. Our study explores an alternative approach to calculating duration empirically and this is done in relation to the banking industry across four countries. Hence, the results of our study will provide further evidence on banks' sensitivity to share prices based on the experience of four countries and the use of a panel cointegration approach to the estimation of duration. As

previously mentioned, this procedure provides the advantage of being able to use more information coming from both time series and cross section data. It also allows for individual effects and therefore makes possible the comparison of duration between banks from different countries. We discuss this procedure in more detail in the ensuing paragraphs.

As also previously stated, Hicks (1939) has shown that duration is also an elasticity measure. As is well known, elasticity is unit free because it can be interpreted in relative terms. It can be shown that elasticity is equal to the estimated slope parameter in a regression equation if the variables are used in logarithmic form. This point can be proved in the following way. Assume that the relationship between the dependent variable (Y) and the independent variable (X) is described by the following exponential function:¹

$$Y_t = AX_t^b e^{u_t} \quad (4)$$

where A is a scale parameter, b is the exponent, e is the natural base and u_t is white noise error term. The mathematical formula for the elasticity of Y with respect to X , denoted by ε_{YX} is the following:

$$\varepsilon_{YX} = \frac{\delta Y_t}{\delta X_t} \frac{X_t}{Y_t} \quad (5)$$

This gives us the following result:

$$\varepsilon_{YX} = \frac{\delta Y_t}{\delta X_t} \frac{X_t}{Y_t} = bAX_t^{b-1} e^{u_t} \frac{X_t}{Y_t} = bAX_t^{b-1} e^{u_t} \frac{X_t}{AX_t^b e^{u_t}} = b \quad (6)$$

¹ Note that the dependent variable is representing the value of the capital asset and the independent variable is representing the interest rate in our case.

Thus, the exponent b is equivalent to the elasticity. In order to estimate the elasticity we need to estimate equation (4). One simple way to do this is to log transform equation (4) as below

$$\ln Y_t = \ln(A X_t^b e^{u_t}) = \ln A + b \ln X_t + u_t \quad (7)$$

Equation (7) can be estimated by OLS that provides a direct measure for the elasticity (duration) in terms of the estimated value for b . However, in this study we will make use of panel data analysis. A corresponding regression model for the panel data is the following:

$$\ln Y_{it} = a_i + b_i \ln X_{it} + \varepsilon_{it} \quad (8)$$

For $i = 1, \dots, N$ and $t = 1, \dots, T$ where N is the number of cross sectional units and T is the time series dimension. We will estimate equation (8) by applying the generalized method of moments (GMM). This method is not sensitive to the existence of autocorrelation and heteroskedasticity. Both of these features might prevail in the panel data set. However, before estimating equation (8) it is of paramount importance to check the time series properties of the underlying panel data in order to avoid spurious and misleading results.

It is now an established fact in the literature that spurious and misleading inference can occur if the time series properties of the underlying data are not carefully taken into account. This is the case because the data generating process for many economic variables is characterised by unit roots. A well-known test statistics for unit roots is the Dickey-Fuller test statistics, which is based on the following regression:

$$\Delta x_t = \gamma x_{t-1} + e_t. \quad (9)$$

The null hypothesis of one unit root is $\gamma = 0$. Dickey-Fuller test statistics has very low power especially in small sample sizes according Shiller and Perron (1985). To improve on the power properties of the Dickey-Fuller test, Levin and Lin (1993) and Im, Pesaran,

and Shin (2003) (IPS hereafter) introduced panel versions of the test. The panel unit-root test is based on the following system:

$$\begin{bmatrix} \Delta x_{1t} \\ \Delta x_{2t} \\ \vdots \\ \Delta x_{Nt} \end{bmatrix} = \begin{bmatrix} \gamma_1 x_{1t-1} \\ \gamma_2 x_{2t-1} \\ \vdots \\ \gamma_N x_{Nt-1} \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \\ \vdots \\ e_{Nt} \end{bmatrix}, \quad (10)$$

Equation (10) combines the time series and cross section dimensions which result in more degrees of freedom. Each error term is assumed to be a white noise random process. The null hypothesis of one panel unit root is $\gamma_1 = \gamma_2 = \dots = \gamma_N = 0$. The panel unit root test that was developed by Levin and Lin (1993) (LL) is based on the following regression:²

$$\Delta x_{it} = \gamma_i x_{it-1} + e_{it}, \text{ for } i = 1, \dots, N \text{ and } t = 1, \dots, T. \quad (11)$$

The number of cross-sections is equal to N and the number of time series observations is equal to T . The combination of these two dimensions results in $N \times T$ degrees of freedom. The panel estimator of the unit root can be defined as the following according to the authors:

$$\sqrt{NT}(\hat{\gamma} - 1) = \frac{\frac{1}{\sqrt{N}} \sum_{i=1}^N \frac{1}{T} \sum_{t=1}^T x_{it-1} e_{it}}{\frac{1}{N} \sum_{i=1}^N \frac{1}{T^2} \sum_{t=1}^T x_{it-1}^2}. \quad (12)$$

The null hypothesis of no panel unit root is tested by applying following t -statistics:

$$t_{\gamma} = \frac{(\hat{\gamma} - 1) \sqrt{\sum_{i=1}^N \sum_{t=1}^T x_{it-1}^2}}{\sqrt{\frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T e_{1t}^2}}. \quad (13)$$

² It should be pointed out that it is possible to add individual constant and trend terms in equation (8) to take into account the effect of the deterministic trend.

It should be mentioned that the Monte Carlo simulation experiments conducted by Levin, Lin and Chu (2002) show that the panel-based unit root test has much better power properties compared to individual unit root test.

The IPS test is more flexible since it allows for a diverse coefficient of unit root. This test is based on an average of the individual Dickey-Fuller tests and it is defined below:

$$\bar{t} = \frac{1}{N} \sum_{i=1}^N t_i, \quad (14)$$

here t_i is the individual t -statistic for testing $H_0: \gamma_i = 0 \forall i, i = 1, \dots, N$. The alternative hypothesis in the IPS test is $\gamma_i < 0$ for $i = N_1 + 1, N_1 + 2, \dots, N$ such that: $\lim_{N \rightarrow \infty} \frac{N_1}{N} = c$, $0 < c \leq 1$. Therefore, this test allows for heterogeneity in the panel. Monte Carlo experiments implemented by Karlsson and Löthgren (2000) also demonstrate that the IPS test has better power properties.

When panel data is used, it is important to check whether the variables have panel unit roots. If they do, then tests for panel cointegration must be performed in order to avoid spurious regression. In this paper, we make use of the procedures developed by Pedroni (1995, 1999, 2004) to test for panel cointegration which are based on the following test statistics:

1. Panel t -Statistic (Non-Parametric):

$$Z_{tN,T} = \left(\tilde{\sigma}_{N,T}^2 \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \left(\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{\lambda}_i \right), \quad (15)$$

2. Panel t -Statistic (Parametric):

$$Z_{tN,T}^* = \left(\tilde{s}_{N,T}^{*2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^{*2} \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^* \Delta \hat{e}_{i,t}^*, \quad (16)$$

3. Group t -Statistic (Non-Parametric):

$$N^{-1/2} \tilde{Z}_{tN,T} = N^{-1/2} \sum_{i=1}^N \left(\hat{\sigma}_i^2 \sum_{t=1}^T \hat{L}_{11}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1/2} \sum_{t=1}^T \left(\hat{e}_{i,t-1} \Delta \hat{e}_{i,t} - \hat{\lambda}_i \right), \quad (17)$$

4. Group t -Statistic (Parametric):

$$N^{-1/2} \tilde{Z}_{tN,T}^* = N^{-1/2} \sum_{i=1}^N \left(\sum_{t=1}^T \tilde{s}_i^{*2} \hat{e}_{i,t-1}^{*2} \right)^{-1/2} \sum_{t=1}^T \hat{e}_{i,t-1}^* \Delta \hat{e}_{i,t}^*, \quad (18)$$

where

$$\hat{\lambda}_i = \frac{1}{T} \sum_{s=1}^{k_i} \left(1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^T \hat{\mu}_{i,t} \hat{\mu}_{i,t-s}, \quad (19)$$

$$\hat{s}_i^2 = \frac{1}{T} \sum_{t=1}^T \hat{\mu}_{i,t}^2, \quad \hat{\sigma}_i^2 = \hat{s}_i^2 + 2\hat{\lambda}_i, \quad (20)$$

$$\tilde{\sigma}_{NT}^2 = \frac{1}{T} \sum_{t=1}^T \hat{L}_{11i}^2 \hat{\sigma}_i^2, \quad \hat{s}_i^{*2} = \frac{1}{T} \sum_{t=1}^T \hat{\mu}_{i,t}^{*2}, \quad (21)$$

and

$$\hat{L}_{11i}^2 = \frac{1}{T} \sum_{t=1}^{k_i} \hat{\eta}_{i,t}^2 + \frac{2}{T} \sum_{t=1}^T \left(1 - \frac{s}{k_i + 1} \right) \sum_{t=s+1}^T \hat{\eta}_{i,t} \hat{\eta}_{i,t-s}. \quad (22)$$

The estimated error terms are obtained from the following regressions:

$$\hat{\varepsilon}_{it} = \hat{\gamma}_i \hat{\varepsilon}_{it-1} + \hat{\mu}_{it}, \quad (23)$$

$$\hat{\varepsilon}_{it} = \hat{\gamma}_i \hat{\varepsilon}_{it-1} + \sum_{k=1}^{K_i} \hat{\gamma}_{ik} \Delta \hat{\varepsilon}_{it-k} + \hat{\mu}_{it}^*, \quad (24)$$

and

$$\Delta \ln Y_{it} = a_0 + a_1 \Delta \ln X_{it} + \hat{\eta}_{it}. \quad (25)$$

Δ is the first difference operator. Note that $\hat{\varepsilon}_{it}$ represents the residuals from the panel model defined by equation (8).

Adjustments are made for each of these test statistics, as suggested by Pedroni, so that each one becomes normally distributed. Thus, the adjusted values for each statistic that we report in this study can be compared to the standard normal distribution values. This applies for both the cointegration and unit root tests.³

³ For more details see Pedroni (1999).

We make use of yearly MSCI banking index data and 3-month interest rate data for Australia, US, UK and Canada during the period 1986-2003. These four countries have developed banking systems in which the banking sector is a major component of their stock markets. They also have fairly independent Central Banks that have the important task of conducting monetary policy. The sensitivity of banks to interest rate would therefore represent important information for the conduct of monetary policy in these countries since banks serve as the channel for which monetary policy decisions are transmitted. Banking, being a major sector, therefore figures heavily in the portfolio allocation decisions of investors and fund managers in each of these countries.

4. Estimation Results

The results for the panel unit roots tests are reported in Table 1. Based on these results, we can conclude that each variable has one panel unit root. Given that the variables are found to be integrated (non-stationary), it is crucial to check whether the variables establish any long-run steady state (cointegration). If the variables that are integrated do not cointegrate, then the regression is spurious. Based on tests for panel cointegration, as suggested by Pedroni, with the results presented in Table 2, we find empirical evidence for panel cointegration between the share price index and the interest rate for all countries in the sample. Two tests reject the null hypothesis of no cointegration at the 1% significance level, one rejects it at 5% level while the remaining test rejects the null at the 10% significance level.

Banking share prices and interest rates, being cointegrated, therefore, share a steady long-run steady relationship. We therefore proceed to estimate the degree of this relationship or elasticity based on regression since the results will not be spurious. The results of our calculation show that, for all countries, bank equity prices react negatively to interest rates. To obtain the country-specific elasticities, we estimated the panel system presented

in equation (8) by making use of GMM method. The estimated elasticities are presented in Table 3 together with their standard errors, t-statistics and the corresponding p-values. It can be seen from Table 3 that all coefficients of elasticity are significant. All coefficients are negative which means that bank share prices react in an opposite direction to interest rates. In absolute terms, the UK has the highest elasticity (1.54), followed by Canada (1.04), Australia (1.00) and then the US (0.60). This indicates that it takes the longest time to recover an investment in bank shares in the UK and shortest in the US. This is supported by the fact that banks in the US are the most profitable while those in the UK are the least (IMF 2005). These results also imply that bank shares in the UK are three times more sensitive to interest rates than those in the US while those in Canada and Australia are almost twice as sensitive. This difference in bank sensitivity to interest rates could be due to the differences in bank profitability as stated earlier. It may be expected that the more profitable banks would be less sensitive to interest rate changes. This differing degree in interest rate sensitivity could also be due to the differences in asset-liability structures and lending spread of the banks in these different countries, as discussed in Section 2 of this paper.

These results could have implications in terms of the ability by banks to act as conduit of monetary policy and for investors. From the results, since the US banks have the highest duration and the UK the lowest, monetary policy changes would have more impact on the banking sector in the UK and least in the US. For investors, this could imply that US banks are more attractive than UK banks.

5. Conclusions

Duration, as initially conceived by McCauley (1938) can be considered as a measure of how quickly an investment can be recovered. As shown by Hicks (1939), it is also an elasticity measure of the relationship between asset prices and interest rates. In this study, we empirically estimate bank equity duration. With the use of new developments in the field of panel cointegration analysis, we investigate the long-run relationship between the share price index for the bank sector and the interest rates for Australia,

Canada, the UK and the US during the period 1986-2003. Several tests for panel unit roots and panel cointegration are performed to avoid the problem of spurious regression in the panel model. The results demonstrate that each variable in the panel is integrated of the first degree. However, the tests for panel cointegration provide empirical support that the variables cointegrate in a panel perspective, which means that bank share prices and interest rates have a stationary long-run relationship.

We found that bank returns are sensitive and negatively related to interest rate changes. Bank shares in the UK have the highest duration, followed by those in Canada, Australia and then the US. Since equity duration implies how fast an investment in shares is recovered, our results indicate that banks in the US are the most attractive for investors while those in the UK are the least. These results also mean that bank shares in the UK are most sensitive to interest rates, followed by those in Canada, Australia and then the US. Thus, interest changes induced by monetary policy would have the greatest impact on banks in the UK and least on banks in the US. The results further imply that US investors on bank shares have the least exposure to interest rate risk, followed by those in Australia, Canada and with those in the UK having the highest. Hence, US bank investors would have the least need for hedging and those in the UK the greatest need.

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Table 1. Test Results for Panel Unit Roots.

H ₀ : I(1), H ₁ : I(0)				H ₀ : I(2), H ₁ : I(1)			
	LL ₁	LL ₂	IPS		LL ₁	LL ₂	IPS
$\ln Y_{it}$	1.551	1.729	2.206	$\ln Y_{it}$	-6.608 ^a	-6.197 ^a	-15.337 ^a
$\ln X_{it}$	0.442	0.124	0.002	$\ln X_{it}$	-4.357 ^a	-4.271 ^a	-6.084 ^a

Notes:

1. Y is representing the share price index of the banking sector in each country and X is representing the interest rate in each country.
2. LL₁ and LL₂ are tests suggested by Levin and Lin (1993). For the first test, the regression is augmented until autocorrelation is removed. The second test adjusts for the effect of potential autocorrelation in the estimation of the parameters.
3. IPS refers to the test recommended by Im et. al. (2003). The notation ^a signifies that the null hypothesis of a panel unit root in the second difference can be rejected at 1% significance level.
4. The adjusted test results are compared to the N(0,1) distribution. Each test is one sided (to the left side of the distribution).

Table 2. Panel Cointegration Test Results for the Share Price Index and the Interest Rate Based on Pedroni Tests.

Test 1	Test 2	Test 3	Test 4
-1.966	-1.337	-1.667	-2.175

Notes: Notice that Test 1 = Panel *t*-Statistic (Non-Parametric), Test 2 = Panel *t*-Statistic (Parametric), Test 3 = Group *t*-Statistic (Non-Parametric), and Test 4 = Group *t*-Statistic (Parametric) as described in the main text. Once again using Pedroni's procedure, we present the adjusted values here that can be compared to the N(0,1). Since the tests are one sided the 1% critical value is -1.96, the 5% value is -1.64 and the 10% critical value is -1.28.

Table 3. The Long-Run Interest Rate Elasticities (Duration) of Bank's Share Prices.

Country	Elasticity	S.E	<i>T</i> -statistics	P-value
Australia	-1.005	0.301	-3.338	0.001
US	-0.604	0.177	-3.412	0.001
UK	-1.548	0.214	-7.228	0.000
Canada	-1.048	0.188	-5.880	0.000

Notes:

1. The elasticity presented in column 2 is the elasticity of bank share prices in each country with respect to interest rates.
2. It should be mentioned that the elasticities are estimated by allowing for individual effects through dummy variables to take into account the scale effect. The GMM method was used to estimate the panel model.

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