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ECONOMY AND THE SHARE MARKET IN AUSTRALIA**

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DISCUSSION PAPER 00.10

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Abstract

In this paper we investigate the effect of financial deregulation on the relationship between the macroeconomy and the share market within the framework of a VAR using quarterly Australian data for three variables – aggregate share prices, real output and the term premium. After an analysis of stationarity and cointegration of the three variables, we specify the VAR in the first differences of the logs of share prices and output and the level of the term premium. Since financial deregulation took place in Australia over an extended period in the 1980s, we do not pre-specify a single break-point but use the Andrews (1993) test for structural stability of the model as a whole. We find a distinct break at 1982(3) which we argue is consistent with the effects of (the early stages of) financial deregulation which began in the late 1970s and continued throughout most of the 1980s. We go on to estimate and simulate the model separately over two sub-samples: 1970(3) – 1982(3) and 1982(4) - 1999(1). We conclude that financial deregulation resulted in greater integration between the share market and the real part of the economy but, paradoxically, weakened the relationship between the share market and other financial markets.

Keywords: share market, VAR model, financial deregulation in Australia

1. Introduction

Both economists and finance specialists are giving increasing attention to the relationship between the share market and the rest of the economy. In many countries there has been remarkable growth of the share market relative to the aggregate economy; in Australia, for example, the ratio of share-market capitalisation to GDP has approximately tripled in the last 25 years – from less than 30% in the mid-1970s to over 80% in the late 1990s.

Not only has the share market increased relative to the real economy, but it appears that the inter-relationship between them has strengthened. It has always been recognised that the share market reflects, at least in the longer term, the goings-on in the rest of the economy but recently there has been widespread recognition that the influence is also in the opposite direction – dramatic events in the share market are likely to have an impact upon the real economy. In many countries there is, e.g., concern about the macroeconomic effects of the possible end to the long bull market.¹ Moreover, with increasingly widespread share ownership,² there is a more direct link between share market events and household behaviour with obvious consequences for the economy as a whole.

Questions concerning the evolving nature and strength of the relationship between the share market and the economy as a whole are therefore important ones in financial economics and it is these which are the focus of this paper. In particular, we first present an estimated model of the relationship between three variables – output, share prices and the term premium – and then go on to an analysis of the stability of the model in the face of sustained financial deregulation in the late 1970s and 1980s. We find strong evidence of a distinct structural break at the end of 1982 which we interpret as being due to financial deregulation. We estimate the model separately for two sub-periods (1970(3)-1982(3), and 1982(4)-1999(1)) and conclude on the basis of model simulations that deregulation resulted in greater integration between the share market and the real part of the economy but that, paradoxically, the relationship between the share market and other financial markets was weaker after the break.

The structure of the paper is as follows. In section 2 we justify our choice of modelling strategy and set out the contribution we make to the existing literature. In

¹ Campbell and Shiller (1998), e.g., characterise the current (1997) state of the US stock market as “extraordinarily bearish” in the light of data for the past 125 years.

² For Australian evidence see Carew(1998), p. 114.

the following section we discuss the data, paying particular attention to the questions of stationarity and cointegration since these properties will determine the specification of the model. In section 4 we present the model and carry out tests of the stability of its coefficients. Model simulations are reported in section 5 and conclusions are drawn in the final section.

2. Modelling Strategy

There are various ways in which the relationship between the share market and the macroeconomy has been modelled in the literature. One approach has been from an asset-pricing perspective in which the Arbitrage Pricing Theory (APT) or some other multi-factor asset-pricing model is used as a framework to address the question of whether risk associated with particular macro variables is reflected in expected asset returns; examples include the original work by Chen, Roll and Ross (1986) who applied the model to the US as did Chen and Jordan (1993); Beenstock and Chan (1988), Clare and Thomas (1994), Cheng (1996) and Antoniou, Garrett and Priestley (1998) all analysed UK data; Ariff and Johnson (1990) used data for Singapore, Martikainen (1991) for Finland and Groenewold and Fraser (1997) for Australia.

A closely-related analysis, based on intertemporal investor optimisation, is that of the consumption-CAPM which concentrates on a single macro influence, the growth of aggregate consumption; see, e.g., Breeden (1979) and Grossman and Shiller (1981).

The direction of influence underlying the asset-pricing literature is the traditional one which is based on the notion that ultimately the share market reflects the fundamental strengths and weaknesses of the aggregate economy so that the direction of influence is from the economy to the share market. A similar focus is found in the literature which explores the response of aggregate share prices to the (expected) inflation rate; early work carried out in this area is by Bodie (1976), Fama and Schwert (1977), Jaffe and Mandelker (1976) and Nelson (1976) whereas more recent applications include those by Balduzzi (1995), Graham (1996), Groenewold, O'Rourke and Thomas (1997) and Siklos and Kwok (1999). Similar studies assess the response of the share market (often, but not always, at an aggregate level) to other macro variables such as those which capture monetary and fiscal policy shocks; e.g. Pearce and Roley (1985), Jain (1988), Aggarwal and Schirm (1992), and Singh (1993).

An alternative, which looks at the influence in the opposite direction, is to analyse the effects of share prices on the macroeconomy or selected macroeconomic variables. A relationship of this nature which has received considerable attention in the financial economics literature is that between share prices and investment (in the sense of capital formation). Studies of this type start with Tobin's q-theory of investment (Tobin, 1969) and also include Fischer and Merton (1984), Morck, Schleifer and Vishny (1990), Blanchard, Rhee and Summers (1993) and Chirinko and Schaller (1996). The question in that literature is whether firms, in making investment decisions should or do pay any heed to share prices or whether share prices are simply a veil which can be dispensed with when making decisions about real variables such as investment.

More recently, essentially empirical models have been used to analyse the relationship between the share market and the real economy. These range from simple single-equation ones of the types used by Chen (1991), Peiro (1996) and Choi, Hauser and Kopecky (1999) to more elaborate models which recognise the two-way relationship between share prices and the economy as a whole. However, unlike the models previously cited, they are not based on any particular theoretical structure but seek simply to capture the empirical regularity between a limited number of variables in a largely pragmatic way.³ The vector auto-regressive (VAR) model has been particularly popular in this area given that it can be used as a framework for formal examination of inter-relationships within a given data set without the need to specify a theoretical framework *a priori*. Once estimated, the model can be used to simulate the effects of shocks in a way that is consistent with the historical patterns in the data by the use of impulse response functions and forecast-error-variance decomposition.

A relatively early application of the VAR model to the analysis of the relationship between share prices and the macroeconomy is by Lee (1992) and more recent ones can be found in Cheung and Ng (1998) and Gjerde and Sættem (1999). Both the Lee and Gjerde and Sættem papers are specified in terms of variables which have been transformed so that they are stationary. It is well known that this involves a potential loss of information and mis-specification if the variables are cointegrated in their levels. The paper by Cheung and Ng explicitly examines whether

³ An alternative approach which is theoretically-constrained is that based on the real-business-cycle (RBC) approach to macroeconomics used by Canova and de Nicolo (1995) for the investigation of the

cointegration is present and finds cointegration between the levels of the variables. It takes the cointegration into account by estimating a variant of the VAR model, viz. the vector error-correction model (VECM).

All three papers using the VAR/VECM approach report impulse response functions, detailing the response of one variable to a shock in another. All three papers find a strong and largely monotonic relationship between output and share prices, although only Lee reports impulse response functions in both directions. He finds a consistently positive response of output to a shock in share prices. Both Lee and Cheung and Ng report a strong and monotonically negative response of share prices to a shock in output (Cheung and Ng provide results for five countries); in contrast, Gjerde and Saettem find that the effect has the opposite sign for Norway. Effects of shocks to other variables such as interest rates and monetary variables are generally mixed and not consistent over time. Thus, while there is strong evidence of a two-way relationship between share prices and output, there is an apparent inconsistency in the literature concerning the sign of influence from the real economy to the share market.

An approach which is closely related to the VAR/VECM procedure is one which is due to Campbell *et al.* – see Campbell and Shiller (1987, 1988) and Campbell and Ammer (1993). More recent applications are by Lee *et al.* – see Lee (1995, 1998), Chung and Lee (1998) and Hess and Lee (1999). While a VAR model is used, the approach differs in at least two ways. First, the VAR model is a constrained one where the constraints are based on a linearised dividend-discount model. It therefore has the advantage of a theoretical structure while at the same time employing the dynamic flexibility of the VAR model. The second difference derives from the first and is that the focus is on the relationship between share prices and other financial variables such as the dividend yield rather than macroeconomic variables such as output. This limits the usefulness of the approach for our purposes.

In this paper we propose to use the VAR/VECM approach, given its flexibility and the absence of any widely-accepted theoretical model of the share-market-economy interrelationship. While the theoretically-restricted Campbell model is attractive, its theoretical restrictions are not applicable to the relationship between share prices and the macroeconomy and we therefore use an unrestricted model.

relationship between real activity and share prices. The extent to which RBC models are empirical is a matter of some controversy. They are better seen as numerical simulation of theoretical models.

We contribute to the literature in various ways. First, we provide evidence of the relationship between share prices and output in a small open economy with relatively mature financial markets and, in so doing, can throw light on the anomalous results reported for Norway by Gjerde and Sættem. Second, following the approach by Cheung and Ng, we specify our model in the light of careful testing for stationarity and cointegration. Third, we report impulse response functions for both directions of the possible relationship between output and share prices. Fourth, we take into account the substantial structural change in Australian financial markets in the 1980s; we test the estimated model for parameter stability using a test by Andrews (1993) which does not require a pre-specified break-point. This is particularly useful because financial-system reform took place over a considerable period and it is not possible to specify a unique break-point with any certainty. We find that the model parameters shifted significantly in the early 1980s and estimate separate models for two sub-samples. We find that the relationship between share prices and output was remarkably stable over the break and that most of the instability detected by the Andrews test involves the interest rate rather than either of the two central variables.

3. The Data

We use a minimal model comprising just three variables: share prices, a real variable (real output) and a financial variable (an interest rate term spread). Even though we are interested primarily in the inter-relationship between output and share prices, we include the financial variable on the basis that other financial markets may be an important channel of transmission of shocks from the real economy to the share market and *vice versa*. The term premium is a useful measure of activity in financial markets since it captures both the long and short ends of the term structure and has been recently used by Peel and Taylor (1998) in analysis of the relationship between financial variables and real output.⁴

⁴ Another reason for using the term premium is that, in contrast to the levels of interest rates, it is stationary over the sample period. The non-stationarity of interest rates is a problem which lead Cheung and Ng to omit them from their model altogether. A non-stationary interest rate seems inconsistent with the observation that in practice interest rates are bounded. Further, as will become clear below, the characteristics of our data would have required us to combine the first difference in the interest rate with first differences in the logs of the share price index and output, a combination of rates of return and growth with differences in a rate of return which seems inconsistent with their theoretical interpretation.

For real output we use GDP (expenditure definition) valued at 1996/97 prices. The data were obtained from Table G.10 of the RBA Bulletin section of the dX database and are available at a quarterly frequency from 1959(3).

Two alternative series were experimented with for share prices. The first is the All Ordinaries (price) index also taken from the RBA Bulletin section of the dX database (Table F.05). It is available on a monthly basis from 1958(1) and was converted to quarterly data to match the frequency of the GDP data by averaging over the three months of each quarter. A drawback with the use of a share price index is that proportional changes in it reflect only capital gains. The alternative is to use a series which includes the effects of dividends. One such series we experimented with was the Total Market Return Index for Australia obtained from Datastream. It is available from 1979(4). Preliminary work with these two series indicated that their time-series properties (such as the autocorrelation function and stationarity) were very similar, confirming experience elsewhere that the inclusion of dividends affects the level of the series but the intertemporal structure is determined largely by price fluctuations. Given that the return index was available for a considerably shorter period, we concentrated on the model using the price index.

Finally, the interest rate spread was between the 10-year T-Bond rate and the 90-day Bank-Accepted Bill rate. Both series were obtained from the RBA Bulletin tables for the dX database, the first from Table F.02 and the second from Table F.01. Data for both variables are monthly and available from 1969(7). They were converted to a quarterly frequency by averaging across the three months of the quarter.

The earliest common starting point for the data is 1969(7) if we use the All Ordinaries price index and we used data until the first quarter of 1999. We use both output and share prices in log form so that first differences have the interpretation of continuously-compounded rates of change and we also use the spread variable in its continuously-compounded form. We now turn to the question of stationarity and, where relevant, cointegration.

Consider the stationarity of the three variables first. We use three tests for stationarity: the augmented Dickey-Fuller (ADF) test, the test due to Phillips and Perron (PP) and the test due to Kwiatkowski, Phillips, Schmidt and Shin (KPSS).⁵

⁵ See Dickey and Fuller (1981), Phillips and Perron (1988) and Kwiatkowski, Phillips, Schmidt and Shin (1992).

All three test statistics depend on the number of lags used in the testing procedure and have, in some applications, been found sensitive to the lag length. Since this was also our experience we begin by checking the appropriate lag length. The use of a Bayesian Information Criterion in the Dickey-Fuller equation suggests that the optimal lag length for the share-price equation is 1, for the output equation is 0 and for the spread equation is 5. This choice was confirmed for the first two variables by inspecting the autocorrelation of the residuals in the equation with the optimal lags and finding that both are free of autocorrelation. In the spread equation autocorrelation was present in the equation with five lags but persisted as the lag length was increased. Hence we proceeded with the test based on five lags but treat the test results with some caution. The test statistics for the chosen lag lengths are presented in Table 1.

Table 1: Stationarity and Cointegration

Variable	ADF		PP		KPSS	
	No Trend	Trend	No Trend	Trend	$\epsilon(\mu)$	$\epsilon(\tau)$
s	-0.1261	-3.0798	-0.0357	-2.6475	5.7067	0.4769
y	-0.0529	0.0239	0.0336	-2.7192	11.7878	0.4728
r	-3.7056	-4.0051	-4.1640	-4.3369	0.3132	0.1213
ds	-8.7124	-6.0398	-8.7124	-8.7282	0.2019	0.1037
dy	-10.4732	-10.4286	-8.7124	-8.7282	0.0660	0.0700
s, y	-3.1089	---	---	---	---	---

Notes: The variables are defined as: s = the log of the share price index, y = log of real GDP, r = spread between the 10-year bond rate and the 90-day bill rate and ds and dy are first differences in s and y.

The 5% critical values for the tests are: ADF(no trend) = -2.86; ADF(trend) = -3.41; PP(no trend) = -2.86; PP(trend) = -3.41; KPSS($\epsilon(\mu)$) = 0.463 and KPSS($\epsilon(\tau)$) = 0.146. For the statistic in the last line of the table the 5% critical value is -3.38.

For the ADF and PP tests the null hypothesis is that there is a unit root in the process (i.e. non-stationarity) while for the KPSS tests the null is that the process is stationary. The results for the levels of the variables (s, y and r) show clearly that both share prices and output are non-stationary and that the spread is stationary. This outcome is robust with respect to the test used and whether there is a trend in the testing equation. Thus r is I(0) and s and y are at least I(1). The next two lines of the table show that the first differences of s and y are stationary showing they are both I(1).

The question therefore arises as to whether s and y are cointegrated. The last line in Table 1 shows the ADF statistic (with 1 lag and no trend) for the test of the stationarity of the residuals from a regression of s on y , i.e. the statistic for the Engle-Granger test for cointegration.⁶ The test suggests that the residuals are non-stationary so that the two variables are not cointegrated.

An alternative tests which is often preferred⁷ to the Engle-Granger two-step procedure is the maximum likelihood-based test of Johansen.⁸ Preliminary investigation showed that the results for this test are also sensitive to lag length in the VAR model within which the tests are carried out. We therefore considered lag length first.

Given that we use quarterly data, we entertained a maximum lag length of 4. Alternative criteria gave somewhat conflicting results: maximisation of the log-likelihood function occurred at four lags, the Akaike criterion suggested three and the Schwartz criterion pointed to zero lags. A formal test of the restriction implied by moving from a VAR(4) to a VAR(3) model is not rejected but the move from a VAR(3) to a VAR(2) model is rejected. Further, both equations in the VAR are free of autocorrelation at lag three. Hence a model with three lags seems an adequate representation of the data for the two variables.

An inspection of the estimated equations show that a trend term is significant in both so that the cointegration test was run with a trend term included. Results are reported in Table 2. In each test, n represents the number of cointegrating vectors so that $n = 0$ implies no cointegration and $n = 1$ implies that the two variables are cointegrated.⁹ Clearly we cannot reject $n = 0$ on the basis of either test and we conclude that there is no evidence of cointegration between share prices and output.

We can therefore work with a straightforward VAR in the first differences of share prices and output and in the level of the interest rate spread. Since all these variables are stationary, we are able to standardise their units of measurement by subtracting the sample mean and dividing by the sample standard deviation. This will not affect the statistical tests but will make the comparison of simulation results

⁶ See Engle and Granger (1987).

⁷ See, e.g., Campbell and Perron (1991) and Davidson and MacKinnon (1993) for a comparison of the tests.

⁸ See Johansen (1988) and Johansen and Juselius (1990).

⁹ While it is common to denote the number of cointegrating vectors by r , we have already used r to represent the term spread and so use n for this parameter.

across variables more informative. We turn to the specification and estimation of such a model in the next section.

Table 2: Johansen Cointegration Tests for (s,y)

Eigenvalue Test			
H_0	H_A	Statistic	5% Critical Value
$n = 0$	$n = 1$	12.4899	18.33
$n \leq 1$	$n = 2$	7.2878	11.54
Trace Test			
H_0	H_A	Statistic	5% Critical Value
$n = 0$	$n \geq 1$	19.7778	23.83
$n \leq 1$	$n = 2$	7.2878	11.54

Note: n represents the number of cointegrating vectors.

4. Specification and Estimation of the Model

Before we estimate the model we must decide on whether to include a trend term and on lag length. The trend term was found to be important in the analysis of stationarity and cointegration in the previous section although the equations used there were specified in levels rather than first differences. We therefore start with a model with four lags of all variables and a trend. In this model the trend term was found to be insignificant in all three equations. Experimentation with shorter lag lengths showed this conclusion to be robust with respect to the number of lags in the model and we therefore do not consider a trend term in the analysis which follows.

Consider now the question of lag length. We started with a maximum lag length of four given that we are working with quarterly data. The various criteria used for deciding lag length gave inconsistent answers: the log of the likelihood function was maximised at a lag of four, the Akaike criterion suggested a lag of three while the Schwarz criterion pointed to a lag length of 1. In formal tests of the restrictions implied by moving from four lags to three and from four lags to two, the restrictions were not rejected but further shortening of the lag length was rejected at the 5% level by a likelihood ratio test. Thus it appears that a model with two lags is required at a minimum. An inspection of the autocorrelations of the residuals of the equations of

the model with two lags indicates an absence of autocorrelation at the 5% level for all three equations. Hence we proceed with a model with two lags and no trend term. Since the variables have zero mean, we omit the intercept term from the model; estimates of the model with an intercept shows it to be insignificant in each of the three equations. The estimated model is reported in Table 3.

Table 3: Estimated Model

Regressor	Equation for ds		Equation for dy		Equation for r	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
ds(-1)	0.2286	2.44	0.0700	0.76	-0.1066	1.83
ds(-2)	-0.0852	0.90	0.1835	1.97	0.1586	2.69
dy(-1)	0.0160	0.17	-0.0089	0.09	-0.0631	1.07
dy(-2)	-0.1406	1.52	-0.1172	1.28	-0.2064	3.58
r(-1)	0.1736	1.22	-0.0131	0.09	0.8643	9.73
r(-2)	-0.0633	0.44	0.1626	1.15	-0.0798	0.90
R ²	0.0909		0.0844		0.6497	
B-G	0.095		0.092		0.072	
RESET	0.060		0.657		0.072	
J-B	0.000		0.014		0.000	
B-P	0.027		0.400		0.027	

Notes: The variables are: ds = first difference in log of the share price, dy = first difference in the log of real output, r = spread between the yield on a 10-year bond and the 90-day bill rate.

The test statistics are: B-G = Breusch-Godfrey test for residual autocorrelation of order four, RESET = Ramsey's test for functional form, J-B = Jarque-Bera test for normality, B-P = Breusch-Pagan test for heteroskedasticity. Prob values are reported for these four tests.

The equations for ds and dy have only limited explanatory power, both having a value for R² of less than 10%. In each case only one regressor is significant, in both cases a lagged value of ds. The lagged values of the term spread are insignificant in both equations. The equation for the spread has better explanatory power with the equation explaining approximately 2/3 of the variation in the dependent variable. In general the diagnostics indicate that each of the equations leaves something to be desired; the share return and spread equations shows evidence of non-normality and heteroskedasticity (both of which are common in financial returns) and the output growth equation showing non-normality in the residuals.

The modest quality of the equations is not surprising given the significant and sustained process of financial-market reform which took place in Australia from the late 1970s and throughout the 1980s. Edey and Gray (1996) point out that under the combined effects of deregulation, market forces and technological developments we have seen “a transformation in the financial system from a relatively closed, oligopolistic structure in the 1950s and 1960s, based predominantly on traditional bank intermediation, to a more open and competitive system offering a much wider variety of services from an array of different providers.” (p.6)¹⁰

While much of the deregulation has focussed on the banks and non-bank financial intermediaries, its impact has undoubtedly been greater – indeed, its wider impact has been the prime motivation for the policy actions driving it. Formal empirical investigation of the effects of the deregulation has been almost exclusively confined to an analysis of the effects on measures of bank risk; relevant references range from early work by Hogan, Sharpe and Volker (1980) and Hogan and Sharpe (1984) to more recent work by Harper and Scheit (1992) and particularly by Faff *et al.* – see Brooks and Faff (1995), Faff and Howard (1997), Brooks, Faff and McKenzie (1997) and Faff and Howard (1999). A notable exception to the focus on banks is the extension of the work in Faff and Brooks (1995) to portfolios other than banks in Brooks and Faff (1997). Broad conclusions for this literature may be taken from Faff and Howard (1999): “Generally, our analysis suggests that the change which occurred in the Australian financial system during the 1980s had important effects on the risks faced by banks and finance companies” (p.99). When the analysis was extended to portfolios for other industries, Brooks and Faff (1997) found that effects on the levels of beta risk differed across industries but that betas became more stable for most industries in the post-deregulation period (p.318).

While the deregulatory measures may have had their initial impact on the financial industries themselves, much of the impetus for the reforms came from the hope that they would improve the working of the economy as a whole. Hence it seems warranted to extend the analysis from the fairly narrow view of the risk of financial firms to the workings of the economy as a whole. In terms of our simple three-variable model, we may expect the dramatic developments in the financial system to have resulted in shifts of the coefficients of the equations as the relationship

¹⁰ A more detailed account of the deregulatory process can be found in Grenville (1991). See also Carew (1998).

between financial markets and the rest of the economy changed and as the relationships between financial markets themselves was altered. We proceed, therefore, to an investigation of the stability of the coefficients of the model.

To use a standard Chow-type test requires the specification of the break-point in advance. Most of the studies cited above have chosen a date early in the 1980s, late 1983 being the most popular since it coincides with the floating of the Australian currency and the removal of very substantial capital account regulations. A weakness of this approach is that tests are conditional on the chosen date, a date which is very difficult to pinpoint given that deregulation took place over an extended time period.

An alternative is to use a more recent test due to Andrews (1993) which allows one to test for structural change in a model with an unknown break-point. The test proceeds as follows. Consider a model with parameter vector β_t for $t = 1, 2, \dots, T$. The null hypothesis is

$$H_0: \beta_t = \beta_0 \text{ for all } t \text{ and some constant } \beta_0.$$

Suppose there is a break point at πT , where $\pi \in (0, 1)$. Then a one-time break at πT is captured by the alternative hypothesis

$$H_1: \beta_t = \beta_1(\pi) \text{ for } t = 1, 2, \dots, \pi T \text{ and } \beta_t = \beta_2(\pi) \text{ for } t = \pi T + 1, \dots, T$$

for some constants $\beta_1(\pi)$ and $\beta_2(\pi)$.

Andrews discusses Wald, LM and LR forms of the test but we concentrate on the LR version. In this case he suggests a test statistic of the form:

$$\sup LR_T(\pi) \text{ for } \pi \in \Pi \text{ where } \Pi \text{ is some sub-set of } [0, 1],$$

where $LR_T(\pi)$ is the LR statistic for a test that there is a structural break at πT . He suggests that Π be set so as to omit approximately 15% of observations at either end of the sample, $\Pi = [0.15, 0.85]$, since the statistic is ill-behaved at the extremes of the sample. Andrews derives the asymptotic distribution of the suggested test statistic and computes its critical values for various degrees of freedom and intervals Π .

The procedure is straightforward to apply to our model since it is linear. We computed the standard LR test statistics for a structural break at each point in the sample (excluding the 15% of the observations at each end of the sample) and compared the maximum value of the test statistic thus computed to asymptotic critical values in Table 1 of the Andrews paper. We used an LR test which permits every parameter to change at the break-point and found a maximum LR statistic of 51.7037 compared to a 5% critical value of 39.55. There is therefore a significant break in the

data and the break occurs at 1982(3) which is consistent with (but slightly earlier than) the break-point chosen a priori by most of the earlier researchers cited above. We therefore feel justified in attributing this break to the effects of financial deregulation. While the maximum value of the test-statistic occurs at 1982(3) we would not want to pinpoint this as the definitive date on which deregulation began to take effect – the statistic was significant (using Andrews' critical values) for all periods after 1982(3) until 1985(3) so that a later break-point is also consistent with the data.

Since we allowed each parameter to change at the break-point, we take the break into account in the estimation of the model by simply estimating it separately over two sub-samples: 1970(1) – 1982(3) and 1982(4) – 1999(1). The results for the first sub-sample are reported in Table 4 and those for the second in Table 5.

Table 4: Estimated Model for 1970(1) – 1982(3)

Regressor	Equation for ds		Equation for dy		Equation for r	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
ds(-1)	0.0719	0.47	-0.1108	0.64	-0.1035	0.89
ds(-2)	0.0065	0.05	0.0885	0.54	0.2058	1.87
dy(-1)	-0.0696	0.53	-0.3237	2.16	-0.0619	0.61
dy(-2)	-0.1274	1.00	-0.3324	2.29	-0.2218	2.27
r(-1)	0.6598	3.40	0.2176	0.98	0.6544	4.38
r(-2)	-0.1137	0.52	0.1494	0.60	-0.1525	0.91
R ²	0.2625		0.1701		0.3587	
B-G	0.018		0.675		0.622	
RESET	0.048		0.340		0.486	
J-B	0.579		0.628		0.000	
B-P	0.898		0.482		0.185	

Notes: The variables are: ds = first difference in log of the share price, dy = first difference in the log of real output, r = spread between the yield on a 10-year bond and the 90-day bank-accepted bill rate. The test statistics are: B-G = Breusch-Godfrey test for residual autocorrelation of order four, RESET = Ramsey's test for functional form, J-B = Jarque-Bera test for normality, B-P = Breusch-Pagan test for heteroskedasticity. In all cases prob values are reported.

There are some noticeable differences between the estimated models for the two sub-periods. In the first place, the degree of explanatory power of the models is generally better than over the full sample period. In the earlier period R² improves for the ds and dy equations while it deteriorates substantially for the r equation. In the

second sub-sample the explanatory power is better for all three equations than it is for the whole period, this being especially the case for the r equation.

Table 5: Estimated Model for 1982(4) – 1999(1)

Regressor	Equation for ds		Equation for dy		Equation for r	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
$ds(-1)$	0.1700	1.40	0.2351	2.39	0.0322	0.60
$ds(-2)$	-0.1611	1.26	0.2028	1.95	0.1149	2.01
$dy(-1)$	0.0486	0.32	0.2256	1.86	-0.0254	0.38
$dy(-2)$	-0.2719	1.94	0.0151	0.13	-0.1122	1.79
$r(-1)$	-0.5149	2.22	-0.1196	0.64	1.2212	11.81
$r(-2)$	0.4557	1.96	0.2904	1.54	-0.3169	3.05
R^2	0.1588		0.3020		0.8736	
B-G	0.258		0.076		0.112	
RESET	0.270		0.185		0.552	
J-B	0.000		0.777		0.386	
B-P	0.968		0.253		0.887	

Notes: The variables are: ds = first difference in log of the share price, dy = first difference in the log of real output, r = spread between the yield on a 10-year bond and the 90-day bank-accepted bill rate. The test statistics are: B-G = Breusch-Godfrey test for residual autocorrelation of order four, RESET = Ramsey's test for functional form, J-B = Jarque-Bera test for normality, B-P = Breusch-Pagan test for heteroskedasticity. In all cases prob values are reported.

In general it is also the case that the diagnostics are better for the sub-samples than they are for period as a whole. The estimated coefficients and their t-ratios show that in the ds equation there has been a shift from influence by lagged r to stronger effects directly from dy and in the dy equation there was a noticeable shift from a predominantly AR process in the earlier period to a strong direct effect of ds in the later period. Both of these results point to a considerable strengthening of the direct links between the real economy and the share market with perhaps a reduction in the importance of interest rates as a channel of influence between these two parts of the economy. However, simple inspection of the coefficients is likely to overlook the more complex dynamic interactions between the two sectors and we turn to an analysis of the impulse response functions for a clearer picture of these.

5. Simulations

The main tool for the analysis of the dynamic properties of our VAR model is the impulse response function (IRF). The IRF is easily derived from the vector moving-average (VMA) form of the model. Write the VAR for an m -variable vector, $\mathbf{x}$, as

$$(1) \quad \mathbf{x}_t = \Phi(L)\mathbf{x}_t + \boldsymbol{\varepsilon}_t, \quad t = 1, 1, \dots, T$$

where $\Phi(L)$ is a p th-order matrix polynomial in the lag operator, L . We assume that $\mathbf{x}_t$ is stationary and that $E(\boldsymbol{\varepsilon}_t \boldsymbol{\varepsilon}_t') = \Sigma$, a positive definite matrix. We have ignored a constant (and other deterministic terms) since they are absent from our particular model. Then the VMA form of the model is obtained from (1) as:

$$(2) \quad \mathbf{x}_t = \mathbf{A}(L)\boldsymbol{\varepsilon}_t,$$

where $\mathbf{A}(L) = (\mathbf{I} - \Phi(L))^{-1}$, an infinite-order matrix polynomial in L . One way of generating IRFs from (2) is to set one of the elements of $\boldsymbol{\varepsilon}_t$ at a non-zero value and all the others at zero and then trace the effects through successive values of $\mathbf{x}_t$. However, this ignores the fact that the elements of $\boldsymbol{\varepsilon}_t$ will generally be correlated so that historically a jump in one of the elements of $\boldsymbol{\varepsilon}_t$ will be associated with changes in other of its elements. A common method of overcoming this difficulty is to re-define the error terms to make them orthogonal so that they can be shocked independently. This is generally achieved by using the Choleski decomposition of the contemporaneous covariance matrix of the errors, Σ . Since Σ is positive definite there exists a lower-triangular matrix (not necessarily unique), $\mathbf{P}$, such that

$$(3) \quad \mathbf{P}\mathbf{P}' = \Sigma$$

The model can then be written in terms of the transformed errors, $\boldsymbol{\xi}_t = \mathbf{P}^{-1}\boldsymbol{\varepsilon}_t$, which are orthogonal. In this case the value of the IRF for the i th element of $\mathbf{x}$ following a shock to the j th error term n periods after the shock is given by

$$(4) \quad \text{IRF}_{ij}(n) = \mathbf{e}_i' \mathbf{A}_n \mathbf{P} \mathbf{e}_j, \quad i, j = 1, 2, \dots, m; n = 0, 1, 2, \dots$$

where $\mathbf{e}_i$ is the i th unit vector and $\mathbf{A}_n$ is the n th matrix in the matrix polynomial $\mathbf{A}(L)$.

While this is a popular procedure, it has the weakness that the orthogonalisation is not unique and the resulting IRFs are not unique but depend on the order in which the variables enter the model. An alternative method, recently suggested by Pesaran and Shin (1998), is to shock a particular error and then to shock all other errors in a way which preserves the historical relationship between them (or

some other assumed correlations). They show that this involves computing the counterpart to (4) as:

$$(5) \quad \text{IRF}_{ij}^G(n) = \sigma_{jj}^{-1} \mathbf{e}_i' \mathbf{A}_n \Sigma \mathbf{e}_j \quad i, j = 1, 2, \dots, m; n = 0, 1, 2, \dots$$

where σ_{jj} is the j th diagonal element of Σ . The advantage of the use of the generalised IRFs is that they are not affected by the ordering of the variables in the model. However, since the shocks in this case are not orthogonal, the IRFs cannot simply be added as they can in the conventional Choleski case. This is not usually a serious weakness since they are generally inspected one at a time. A more serious drawback of the use of the generalised procedure of Pesaran and Shin is that the forecast-error-variance decompositions (FEVDs) associated with it do not sum to unity as they do when the Choleski decomposition is used. For our purposes this is not particularly serious since we concentrate on the IRFs.

The IRFs for our model are graphed in Figures 1, 2 and 3. We report IRFs only for the model estimated over the two sub-periods since it is clear that the model estimated over the full sample is mis-specified. In each case we graph the IRF before and after the break in the same diagram and concentrate on a comparison of the dynamic behaviour before and after the break to assess the effect of financial-market reform on the dynamic interaction between the share market and the real economy. In all cases the IRF before the break is depicted by a solid line and the IRF after the break with a broken line. We show results only for an horizon of 10 quarters since in almost all cases the action is over by the time ten quarters have elapsed. Figure 1 pictures the effects of a shock to ds on ds itself, on dy and on r (Figures 1(a), 1(b) and 1(c) respectively). Figures 2 and 3 are similar but picture the effects of shocks to dy and r respectively.

[Figure 1 near here]

Consider the effects of a shock to ds first as shown in Figure 1. The effect of a share-market shock on the share market itself dies out very quickly and there is little change to the IRF over the break although there is more correction in the second sub-sample than there is in the first which may be reflected in greater volatility post-deregulation. The effect of the shock on output is predominantly positive and the effect is markedly stronger after the break than it is before.

Of the three VAR/VECM papers cited above, only Lee (1992) reports the effect of a share-market shock on output and he also reports a positive sign. He offers the explanation that it reflects the share market's leading the economy in the sense that share prices anticipate changes in real output. The positive sign may also be explained by the wealth effect – a boost to share prices has a positive wealth effect on consumption which boosts output through conventional aggregate demand channels. This explanation is consistent with the stronger post-break effect which would then reflect the increasing importance of share-holding over the sample. It seems at first sight inconsistent with a substitution story – a rise in returns to shares causes a substitution in portfolios out of other assets into shares which should raise returns to other assets and therefore depress output through the standard interest-rate effects on aggregate demand. The first link in the chain is borne out by the IRF for the effect of a ds -shock on r (it is strongly positive) but rather than a fall in dy as the macro story requires, we observe a rise in dy (again borne out by the effect of an r -shock on dy in Figure 3(b)). However, it should be recalled that the interest-rate variable in the model is the term spread so a rise in r is consistent with a fall in the short rate which may boost interest-sensitive aggregate demand components. The effects of ds on r have already been mentioned in passing and the pattern of the effect also appears stable over the break-point.

[Figure 2 near here]

Consider now the effects of an output shock as pictured in Figure 2. The effect on output itself dies out very quickly, more so after 1982(3). The effect of a dy -shock on ds is strongly negative and the pattern of the effect does not differ across the break-point although, as with the reverse effect, the amplitude of the fluctuations is somewhat larger after the break. This reinforces our earlier conclusion that the relationship between the real economy and the share market have strengthened after the break attributed to financial deregulation – the share market has become more integrated into the economy as a whole in this sense.

The negative effect of output shocks on dy is consistent with that reported both by Lee (1992) for the US and by Cheung and Ng (1998) for a set of five

countries¹¹ but inconsistent with the positive sign found by Gjerde and Sættem (1999) for Norway. None of these authors offers any explanation for the sign. At first sight the positive sign seems more plausible in a standard macro framework – a demand-driven rise in output raises interest rates and this drives up returns to shares through a portfolio-substitution effect. However, a supply-driven output expansion will have the opposite effect since, *cet. par.*, a rise in aggregate supply, given aggregate demand, will depress prices (or reduce inflation in an inflationary world), expanding real money and therefore depressing interest rates and thus share returns through the substitution effect. It is possible, therefore, that the different effect reported for Norway is not linked to the nature of the Norwegian economy (Gjerde and Sættem point out that it is an small open economy with relatively under-developed financial markets) but may simply reflect a different balance between supply- and demand-driven output changes over the sample period.

[Figure 3 near here]

Turning now to the effects of interest-rate shocks pictured in Figure 3, we observe dramatic shifts in the overall shapes of the IRFs. The effects of a term-spread shock on share prices before the break is almost the mirror image of the post-break effects – in both cases the effect dies out within a year but before 1982 it is predominantly positive whereas after that date it is almost completely negative. The same marked change in behaviour is true of the response of output to a shock to r with the added characteristic that the effect dies out much more slowly after the break than before, a characteristic shared with the response of r itself to a shock to the r -equation. It is possible that this reversal can be rationalised in terms of shifts from demand to supply shocks being the main driving force of the change in the term premium since they will have different effects on output.¹² An alternative and not mutually exclusive explanation for the instability of the interest-rate effects in our model is that it reflects the widely-observed instability of the demand for money over the same period and is, no doubt, related to the same driving forces, an important component of which is financial deregulation and technological change in financial markets.¹³

¹¹ Canada, Germany, Italy, Japan and the US.

¹² In a paper which uses the Blanchard and Quah (1989) scheme to distinguish between demand and supply shocks, Peel and Taylor (1998) show that for the US and UK over the period 1957-1994 the term premium affected output mainly through demand channels.

¹³ For a survey of empirical literature on the demand for money see, e.g. Goldfeld (1992).

Consider, finally, the FEVDs reported in Table 6. As with the IRFs, we use the generalised versions based on Pesaran and Shin (1998). In general the components reported do not add to unity but in most cases the discrepancy is small. The relative magnitudes of the components of the FEVDs are comparable to those reported elsewhere – see, e.g., the papers by Lee (1992) and Gjerde and Sættem (1999) discussed earlier. Moreover, the results reported show the standard feature that by far the largest proportion of the variance is explained by the innovation in the variable's own equation.

Table 6: Forecast-Error-Variance Decompositions

Error-variance of:		Error-variance due to innovations in:					
		ds		dy		r	
		pre- 1982(3)	post- 1982(3)	pre- 1982(3)	post- 1982(3)	pre- 1982(3)	post- 1982(3)
ds,	n = 1	0.8162	0.9498	0.0054	0.0018	0.2406	0.0475
	n = 10	0.7953	0.9058	0.0244	0.0377	0.2584	0.0572
dy,	n = 1	0.0072	0.0745	0.9791	0.9237	0.0420	0.0028
	n = 10	0.0205	0.1639	0.9619	0.7849	0.0459	0.0526
r,	n = 1	0.0350	0.0051	0.0332	0.0007	0.9840	0.9971
	n = 10	0.0962	0.0466	0.0984	0.1032	0.8572	0.8565

The decomposition of the ds variance shows that there was a marked shift over the break-point – in the first part of the sample approximately 24% of the ds-variance was explained in the long run by innovations in the interest-rate equation indicating a strong link between these two financial markets. However, after 1982 this proportion fell to about 5% with a slight increase in the share of the dy-variance but most being taken up by the innovation in the ds-equation itself.

When we move to the variance of forecasts of dy, we find that, by way of contrast, the influence of the ds-innovation increased considerably after 1982 pointing to much greater direct effects of the share market on the real economy as we have found from our analysis of the IRFs. The influence of the r-innovation was relatively weak both before and after the break-point.

The FEVDs for r show that before 1982, the term spread was influenced by both ds and dy innovations but that after 1982, the relationship with the share market was dramatically weakened in both directions.

In conclusion, we find that there are inter-relationships between components of the model which are consistent in magnitude with those reported for other countries and time periods. We find also that there has been a marked shift in the nature of these inter-relationships in 1982. While the evidence is not altogether clear-cut, there is a strong suggestion that the share market has become more closely related to the real output market (primarily in the direction from the share market to output) and clearly less tightly connected to the fixed-interest security markets (in both directions). Thus financial de-regulation seems to have strengthened the influence of the share market on the real economy but weakened the inter-relationships between the share market and other financial markets.

6. Conclusions

In this paper we have been concerned with the relationship between the share market and the rest of the economy and with the way in which this relationship may be affected by the deregulation of financial markets which has occurred in Australia over the 1980s. We carried out our investigation within the framework of a minimal empirical model which is able to account for two-way influences between all the variables. We chose to use a model of the VAR/VECM-type in three variables – aggregate share prices, real output and the term premium. A preliminary analysis of stationarity and cointegration pointed to a model specified in terms of the first differences in the logs of output and share prices and the level of the term premium.

The model performed only poorly when estimated over the entire sample period from 1970-1999. In particular, it failed a test for structural stability – the application of a test due to Andrews (1993) showed a strong break at 1982(3) which we argued is consistent with the effects of financial deregulation in Australia during the 1980s and also consistent with the break-point chosen in much of the existing literature on the effects of the deregulation of the Australian financial markets. We therefore estimated and simulated the model separately over two sub-samples: 1970(3) – 1982(3) and 1982(4) - 1999(1). The model performed markedly better over the separate sub-samples.

The simulation results showed a positive response of output to share-market shocks and a negative response in the opposite direction. We provided a rationalisation for each of these effects within a standard macroeconomic framework. Both results were robust over the break at 1982(3) although the effects were stronger after the break suggesting that financial deregulation strengthens the inter-relationship between the real economy and the share market but does not change its direction. The opposite was the case for most of the simulations involving the term spread – the direction of the effects was often reversed over the break-point of 1982(3) and the relationship between share prices and the term spread weakened over the period.

Thus financial deregulation resulted in greater integration between the share market and the real part of the economy but, paradoxically, weakened the relationship between the share market and other financial markets (captured by the term premium).

References

- Aggarwal, R and D. C. Schirm (1992), "Balance of Trade Announcements and Asset Prices: Influence on Equity Prices, Exchange Rates and Interest Rates", Journal of International Money and Finance, 11, 80-95.
- Andrews, D. W. K. (1993), "Tests for Parameter Instability and Structural Change with Unknown Change Point", Econometrica, 61, 821-856.
- Antoniou, A., I. Garrett and R. Priestley (1998), "Macroeconomic Variables as Common Pervasive Risk Factors and the Empirical Content of the Arbitrage Pricing Theory", Journal of Empirical Finance, 5, 221-240.
- Ariff, M. and L. W. Johnson (1990), "Ex Ante Risk Premia on Macroeconomic Factors in the Pricing of Stocks: An Analysis Using Arbitrage Pricing Theory", Chapter 18 in M. Ariff and L.W. Johnson, Securities Markets and Stock Pricing, Longman, Singapore, 194-204.
- Balduzzi, P. (1995), "Stock Returns, Inflation and the 'Proxy Hypothesis': A New Look at the Data", Economics Letters, 48, 47-53.
- Beenstock, M. and K. Chan (1988), "Economic Forces in the London Stock Market", Oxford Bulletin of Economics and Statistics, 50, 27-39.
- Blanchard, O. J. and D. Quah (1989), "The Dynamic Effects of Demand and Supply Disturbances", American Economic Review, 79, 655-673.
- Blanchard, O. J., C. Rhee and L. Summers (1993), "The Stock Market, Profit and Investment" Quarterly Journal of Economics, 108, 115-136.
- Bodie Z., (1976), "Common Stocks As a Hedge Against Inflation", Journal of Finance, 31, pp. 459-470.
- Boudoukh, J. and M. Richardson, (1993), "Stock Returns and Inflation: A Long-Horizon Perspective", American Economic Review, 83, 1346-1355.
- Breeden, D. (1979), "An Intertemporal Asset Pricing Model with Stochastic Investment and Consumption Opportunities", Journal of Financial Economics, 7, 265-296.
- Brooks, R. D. and R. W. Faff (1995), "Financial Market Deregulation and Bank Risk: Testing for Beta Instability", Australian Economic Papers, 34, 180-199.
- Brooks, R. D. and R. W. Faff (1997), "Financial Deregulation and Relative Risk of Australian Industry", Australian Economic Papers, 36, 308-320.
- Brooks, R. D., R. W. Faff and M. D. McKenzie (1997), "Bivariate GARCH Estimation of Beta Risk in the Australian Banking Industry", Accountability and Performance, 3, 81-101.

- Campbell, J. Y. and J. Ammer (1993), "What Moves the Stock and Bond Markets? A Variance Decomposition for Long-Term Asset Returns", Journal of Finance, 48, 3-37.
- Campbell, J.Y. and P. Perron (1991), "Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots", in O.J. Blanchard and S. Fischer (eds.), Macroeconomics Annual 1991, Cambridge: MIT Press, 141-201.
- Campbell, J. Y. and R. J. Shiller (1987), "Cointegration and Tests of Present Value Models", Journal of Political Economy, 95, 1062-1088.
- Campbell, J. Y. and R. J. Shiller (1988), "Stock Prices, Earnings and Expected Dividends", Journal of Finance, 43, 661-676.
- Campbell, J. Y. and R. J. Shiller (1998), "Valuation Ratios and the Long-Run Stock Market Outlook", Journal of Portfolio Management, 24, 11-26.
- Carew, E. (1998), Fast Money 4, Allen and Unwin, Sydney.
- Canova, F. and G. De Nicrolo (1995), "Stock Returns and Real Activity: A Structural Approach", European Economic Review, 39, 981-1015.
- Chen, N. (1991), "Financial Investment Opportunities and the Macroeconomy", Journal of Finance, 46, 529-554.
- Chen, N., R. Roll and S. A. Ross (1986), "Economic Forces and the Stock Market", Journal of Business, 59, 383-403.
- Chen, S. J. and B. D. Jordan (1993), "Some Empirical Tests of the Arbitrage Pricing Theory: Macro Variables vs. Derived Factors", Journal of Banking and Finance, 17, 65-89.
- Cheng, A. C. S. (1996), "The UK Stock Market and Economic Factors: A New Approach", Journal of Business Finance and Accounting, 22, 129-142.
- Cheung, Y.-W. and L. K. Ng (1998), "International Evidence on the Stock Market and Aggregate Economic Activity", Journal of Empirical Finance, 5, 281-296.
- Chirinko, R. S. and H. Schaller (1996), "Bubbles, Fundamentals and Investment: A Multiple Equation Testing Strategy", Journal of Monetary Economics, 38, 47-76.
- Choi, J. J., S. Hauser and K. J. Kopecky (1999), "Does the Stock Market Predict Real Activity? Time Series Evidence from the G-7 Countries", Journal of Banking and Finance, 23, 1771-1792.
- Chung, H. and B.-S. Lee (1998), "Fundamental and Nonfundamental Components in Stock Prices of Pacific-Rim Countries", Pacific-Basin Journal of Finance, 6, 321-346.

- Clare, A. D. and S. N. Thomas (1994), "Macroeconomic Factors, the APT and the UK Stock Market", Journal of Business Finance and Accounting, 21, 309-330.
- Dickey, D. A. and W. A Fuller (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", Econometrica, 49, 1057-1072.
- Edey, M. and B. Gray (1996), "The Evolving Structure of the Australian Financial System", in M. Edey (ed.), The Future of the Financial System, Reserve Bank of Australia, Sydney, 6-44.
- Engle, R. F. and C. W. J. Granger, (1987), "Cointegration and Error Correction: Representation, Estimation and Testing", Econometrica, 55, 251-276.
- Faff, R. W. and P. F. Howard (1997), "Bank Exposure to Interest-Rate Risk: The Case of the Australian Banking Industry", Applied Economics Letters, 4, 737-739.
- Faff, R. W. and P. F. Howard (1999), "Interest Rate Risk of Australian Financial Sector Companies in a Period of Regulatory Change", Pacific-Basin Finance Journal, 7, 83-101.
- Fama, E. F. and G. W. Schwert, (1977), "Asset Returns and Inflation", Journal of Financial Economics, 5, 115-146.
- Fischer, S. and R. C. Merton (1984), "Macroeconomics and Finance: The Role of the Stock Market", Carnegie-Rochester Conference Series on Public Policy, 21, 57-108.
- Gjerde, O. and F. Sætem (1999), "Causal Relations Among Stock Returns and Macroeconomic Variables in a Small, Open Economy", Journal of International Financial Markets, Institutions and Money, 9, 61-47.
- Goldfeld, S. M. (1992), "Demand for Money: Empirical Studies" in P. Newman, M. Milgate and J. Eatwell (eds.), The New Palgrave Dictionary of Money and Finance, Vol. 1, 618-624.
- Graham, F. C. (1996), "Inflation, Real Stock Returns and Monetary Policy", Applied Financial Economics, 6, 29-35.
- Grenville, S. (1991), "The Evolution of Financial Deregulation", in I. MacFarlane (ed.), The Deregulation of Financial Intermediaries, Reserve Bank of Australia, Sydney, 3-35.
- Groenewold, N. and P. Fraser (1997), "Share Prices and Macroeconomic Factors" Journal of Business Finance and Accounting, 24, 1367-1383.
- Groenewold, N., G. O'Rourke and S. Thomas (1997), "Stock Returns and Inflation: A Macro Analysis", Applied Financial Economics, 7, 127-136.

- Grossman, S. and R. Shiller (1981), "The Determinants of the Variability of Stock Market Prices", American Economic Review, 71, 222-227.
- Gultekin, N.B. (1983), "Stock Market Returns and Inflation: Evidence from Other Countries", Journal of Finance, 38, 49-65.
- Harper, I. R. and T. Scheit (1992), "The Effects of Financial Market Deregulation on Bank Risk and Profitability", Australian Economic Papers, 31, 260-271.
- Hess, P. J. and B.-S. Lee (1999), "Stock Returns and Inflation with Supply and Demand Disturbances", The Review of Financial Studies, 12, 1203-1218.
- Hogan, W. P. and I. G. Sharpe (1984), "Regulation, Risk and the Pricing of Australian Bank Shares", Economic Record, 60, 34-44.
- Hogan, W. P., I. G. Sharpe and P. A. Volker (1980), "Risk and Regulation: An Empirical Test of the Relationship", Economics Letters, 6, 373-379.
- Jaffe, J. and G. Mandelker, (1976), "The 'Fisher Effect' for Risky Assets: An Empirical Investigation", Journal of Finance, 31, 447-458.
- Jain, P. C. (1988), "Response of Hourly Stock Prices and Trading Volume to Economic News", Journal of Business, 61, 219-231.
- Johansen, S. (1988), "Statistical Analysis of Cointegration Vectors", Journal of Economic Dynamics and Control, 12, 231-254.
- Johansen, S. and K. Juselius (1990), "Maximum Likelihood Estimation and Inference on Cointegration - with Applications to the Demand for Money", Oxford Bulletin of Economics and Statistics, 52, 169-210.
- Kwiatkowski, D., P. C. B. Phillips, P. Schmidt and Y. Shin (1992), "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure Are We That Economic Time Series Have a Unit Root?", Journal of Econometrics, 54, 159-78.
- Lee, B.-S. (1992), "Causal Relations Among Stock Returns, Interest Rates, Real Activity and Inflation", Journal of Finance, 47, 1591-1603.
- Lee, B.-S. (1995), "The Response of Stock Prices to Permanent and Temporary Dividend Shocks", Journal of Financial and Quantitative Analysis, 30, 1-22.
- Lee, B.-S. (1998), "Permanent, Temporary and Non-Fundamental Components of Stock Prices", Journal of Financial and Quantitative Analysis, 33, 1-32.
- Martikainen, T. (1991), "On the Significance of the Economic Determinants of Systematic Risk: Empirical Evidence with Finnish Data", Applied Financial Economics, 1, 97-104.

- Morck, R., A. Schleifer and R. W. Vishny (1990), "The Stock Market and Investment: Is the Market a Sideshow?", Brookings Papers on Economic Activity 2, 157-215.
- Nelson, C. R., (1976), "Inflation and Rates of Return on Common Stock", Journal of Finance, 31, 471-483.
- Pearce, D. K. and V. V. Roley (1985), "Stock Prices and Economic News", Journal of Business, 58, 49-67.
- Peel, D. A. and M. P. Taylor (1998), "The Slope of the Yield Curve and Real Economic Activity: Tracing the Transmission Mechanism", Economics Letters, 59, 353-360.
- Peiro, A. (1996), "Stock Prices, Production and Interest Rates: Comparison of Three European Countries with the USA", Empirical Economics, 21, 221-234.
- Pesaran, M. H. and Y. Shin (1998), "Generalised Impulse Response Analysis in Linear Multivariate Models", Economics Letters, 58, 17-29.
- Phillips, P. C. B. and P. Perron (1988), "Testing for a Unit Root in Time Series Regression", Biometrika, 75, 335-346.
- Siklos, P. and B. Kwok (1999), "Stock Returns and Inflation: A New Test of Competing Hypotheses", Applied Financial Economics, 9, 567-581.
- Singh, R. A. (1993), "Response of Stock Prices to Money Supply Announcements: Australian Evidence", Accounting and Finance, 33, 43-59.
- Tobin, J. (1969), "A General Equilibrium Approach to Monetary Theory", Journal of Money, Credit and Banking, 1, 15-29.

Figure 1: Effects of a shock to ds

Figure 2: Effects of a shock to dy

Figure 3: Effects of a shock to r