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Abstract:
This paper analyses the relationship between stock prices and the Chinese macro economy measured by the level of GDP. There are many possible channels of influence between these two variables, channels which may operate in either direction. There are also many theories relevant to these interrelationships. Rather than explicitly testing theories, we focus on the empirical nature of this relationship which we analyse in the context of a VAR/VEC model which allows for two-way influences but is agnostic about the particular theoretical underpinnings. We apply tests for stationarity and cointegration and find that there is a long-run, cointegrating relationship between stock prices and GDP. We estimate a VEC model and use it to analyse both short-run and long-run causality as well as to generate impulse response functions (IRFs). We find that there is strong evidence of long-run causality from the economy to the stock market but not *vice versa*. We also find modest but weaker evidence of a similar short-run effect. These are borne out by the IRFs which show a small and weak link from the stock market to the economy but a stronger and much more substantial effect in the opposite direction. We rationalise our results in terms of the relatively small size of China's stock market.

I Introduction

Recent sharp changes in stock prices prompt a host of questions about the relationship between the stock market and the rest of the macroeconomy. At a broad level there are questions about the benefits of a deregulated financial system for the health of the economy while questions with a narrower focus are also common, such as: do the fluctuations in stock prices reflect economic factors or are they simply bubbles driven by (irrational) investor sentiment? On the other hand, do the stock price movements spill over into the rest of the economy, via consumption, investment or some other channel?

Given the importance of the stock market in the financial system of most countries, it is not surprising that all of these questions have regularly exercised the minds of policy-makers and have been the subject of a substantial amount of empirical research. The broad question of the relationship between financial development and economic development in general is the subject of a rapidly-growing literature which has been recently surveyed in a wide-ranging paper by Levine (2005). Measures of financial development in this literature include more than just those relating to the stock market and range from assets and liabilities of banks and non-bank financial intermediaries to the size of the bond market relative to the economy as a whole, as well as stock market capitalisation and turnover variables. The focus of the research in this area is long-run and typically uses large cross-country data bases with variables measured as multi-year averages to capture the long-term nature of the growth process although, more recently, panel data have also been used to enable the analysis of questions of causality in the finance-growth nexus. Levine draws the overall conclusion that there is a positive connection between measures of financial

development and economic growth and that stock market development makes a significant contribution to this effect.

In addition to questions of long-run growth, the stock-market-macroeconomy connection has also been analysed from a short-run perspective, focussing on the relationship between stock prices and macroeconomic variables such as GDP, consumption, investment, inflation, exchange rates and monetary policy measures. Here relationships may run in both directions, from the macroeconomy to the stock market and *vice versa*. While the evidence points to a positive effect of stock prices on output there is mixed evidence on the sign of the effect in the opposite direction.

The present paper reports on the analysis of the relationship between stock prices and the macroeconomy in China for the period since the establishment of the Chinese stock market in the early 1990s. Our focus is on the short-run interaction between stock prices and the macroeconomy, in contrast to the long-run emphasis of the finance and growth literature. Moreover, we confine our attention to output as the main measure of the macroeconomy.

Our motivation for this research is three-fold. First, the relationship between output and stock prices is not clear; in particular, negative effects of output shocks on stock prices have been reported in several papers such as Lee (1992) for the US, Cheung and Ng (1998) for a set of five countries and Groenewold (2003) for Australia. This is in contrast to other findings such as those by Gjerde and Saettem (1999) and seems counterintuitive although Groenewold suggests an explanation which distinguishes demand- from supply-driven output shocks. Analysis of this relationship for different countries will provide more information on this important connection.

Secondly, China's stock market is relatively small although it is developing rapidly and an analysis of this case will balance the predominance of developed economy research.

Thirdly, little is known about the output-stock-price nexus in China. There is a very limited Chinese-language literature and, to our knowledge, only two English-language papers address this issue, namely Zhao (1999) and Liu and Sinclair (2008), only the latter of which throws any light on the issue. Given the growing importance of China in the world economy and in the international financial system, the relationship between stock prices and macroeconomic variables is an important issue in its own right and deserves a more thorough investigation.

The remainder of the paper is structured as follows. The next section presents background material by way of a review of the relevant literature while section III presents background material on the Chinese stock market and briefly puts it in the perspective of the Chinese financial system. In section IV we describe our data and report the results of tests for stationarity. Tests for cointegration, the estimated VECM and the associated tests of causality (both short- and long-run) and impulse response functions which we use as the main instruments to address the issue of this paper are reported in section V. We present conclusions in section VI.

II Literature

There are various ways in which the short-run relationship between the stock 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 Kim and Wu (1987) and Chen and Jordan (1993). There have been numerous applications to other countries' stock prices such as: Beenstock and Chan (1988), Clare and Thomas (1994), Cheng (1996), Antoniou, Garrett and Priestley (1998) and Gunsel and Cukur (2007) for the UK; Entorf and Jamin (2007) for Germany; Tsuji (2007) for Japan; Ariff and Johnson (1990) for Singapore; Martikainen (1991) for Finland; Groenewold and Fraser (1997) for Australia; Mateev and Videv (2008) for Bulgaria and Ihsan *et al.* (2007) for Pakistan.

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). Applications/tests have been reported in Breeden *at al.* (1989), Kocherlakota (1997), Cashin and McDermott (1998) and Chen (2003).

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 in the spirit of the Fisher effect. Early work carried out in this area is by Bodie (1976), Fama and Schwert (1977), Jaffe and Mandelker (1976), Nelson (1976) and Gultekin (1983) whereas more recent applications include those by Boudoukh and Richardson (1993), Balduzzi (1995), Graham (1996), Groenewold, O'Rourke and Thomas (1997), Siklos and Kwok (1999), Crosby (2001) and Boucher (2006).

Related 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), Singh (1993), Thorbecke (1997), Cassola and Morana (2004), Wong et al. (2006) and Ioannidis and Kontonikas (2008).

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 economist's sense of capital formation). Studies of this type start with Tobin's q-theory of investment (Tobin, 1969) and include Fischer and Merton (1984), Morck, Schleifer and Vishny (1990), Blanchard, Rhee and Summers (1993), Chirinko and Schaller (1996, 2001), Baker *et al.* (2003) and Gilchrist *et al.* (2004). 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.

Another route through which stock prices have been seen to influence the real economy is via consumption, the most common theoretical basis being the wealth effect in the consumption function although other channels such as increased uncertainty (Romer, 1990), signalling effects (Poterba and Samwick,1995, Poterba, 2000), realised and unrealised wealth effects, liquidity-constraint effect and a stock-option value effect (Ludwig and Sløk, 2002) have also been suggested. Empirical work on the stock-price-consumption relationship has included cross-section, time-series and mixed (panel) studies; examples are Mankiw and Zeldes (1991), Parker (1999), Starr-McCluer (2002), Ludvigson and Steindel (1999), Shirvani and Wilbratte

(2000, 2002), Case *et al.* (2001), Edison and Sløk (2002), Bertaut (2002), Lettau and Ludvigson (2004) and Sousa (2008).

In addition to the theoretically-informed analysis reviewed above, essentially atheoretical empirical models have also been used to analyse the relationship between the share market and the macroeconomy. These range from simple single-equation ones of the types used by Chen (1991), Peiro (1996), Choi, Hauser and Kopecky (1999) and Ioannidis and Kontonikias (2008) 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.[1] The vector auto-regressive (VAR) and vector error-correction (VEC) models have been particularly popular in this area, given that they 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 decompositions.

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 Thorbecke (1997), Cheung and Ng (1998) and Gjerde and Saettem (1999), Cassola and Morana (2004), McMillan (2005), Phylatkis and Panazzolo (2005), Wong *et al.* (2006) and Pan *et al.* (2007).

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[1] 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 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.

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 derived from 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.

We conclude this section with a brief account of the limited literature on the output-stock-price relationship in China. In the Chinese language literature there are several papers which deal with our topic although there appears to be some confusion between output and growth, with several papers claiming to be an analysis of stock prices and economic growth but actually analysing the relationship between stock prices and GDP (often both in levels) so that they are directly relevant to the work reported in this paper. Examples are: Ran, Zhang and Wu (2003), Ran, Hu and Wang (2005), Liang and Teng (2005) and Fan (2006). These papers generally test for stationarity and cointegration in the (logs of) stock prices and macro variables, principally output, and then go on to test for causality between them. Interesting variations are recent papers by Wei and Yong (2007) and Han, Zhang and Wu (2008), the latter of which focuses on inflation and stock prices and decomposes inflation shocks into supply and demand-driven ones which, it is found, have different effects on stock prices.

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Finally, two English-language papers are also related to our work, the first by Zhao (1999) and the second by Liu and Sinclair (2008). The Zhao paper uses a single equation framework to regress the rate of change of output (so growth rather than output) on the rate of growth of stock prices. A distinction is made between total growth and unexpected growth and regressors are entered contemporaneously. The finding most relevant to our work is that total growth has a negative and significant effect on stock returns but the unexpected component of growth has a positive and significant effect. The reverse effect from stock prices to output growth is not tested.

Interestingly, the Liu and Sinclair paper purports to analyse the question of the relationship between the stock market and economic growth (for Hong Kong and Taiwan as well as for mainland China) but, in fact, analyses the relationship between the log of stock prices and the log of output in a VECM framework, thus being more closely related to our work than the Zhao paper is. They find short-run causality running from stock prices to output but not *vice versa* but claim that output affects stock prices in the long run, although they do not present test results for this hypothesis.

The existing literature on the relationship between the stock market and the economy as a whole in China is thus very limited and contradictory and considerably more through-going analysis is necessary before the relationship is well understood. We intend to contribute to such analysis.

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 directly applicable to the relationship between share prices and the macroeconomy and we therefore use an unrestricted

model. Before turning to the modelling framework and the empirical analysis, we digress briefly to present some basic information about the Chinese stock market and the financial system in general to provide background to the interpretation of our results.

III The Chinese stock market and the Chinese financial system

The Chinese stock market consists of two exchanges, the Shanghai Securities Exchange (SHSE) and the Shenzhen Securities Exchange (SZSE). The SHSE was opened in December, 1990 and the SZSE in February, 1991. Since 1998, the market has been supervised by the Chinese Securities Supervision Commission, before which it was regulated by a State Council committee. While it has been subject to many complicated regulations, including price limits from time to time, the trend is towards cautious deregulation.

An interesting feature of the market for the first two decades of its existence is various types of shares. The two main types are A and B shares.[2] A shares are denominated in the local currency (Renminbi or RMB) and are traded by domestic residents and institutions – foreign individuals and institutions are not permitted to buy and sell A shares. B shares are denominated in US dollars on the Shanghai Exchange and Hong Kong dollars on the Shenzhen Exchange. They were originally intended for trading by foreign investors but the restriction that only offshore individuals and institutions are permitted to trade in B shares was lifted in 2001, permitting domestic residents to trade in them but only in foreign currency. In

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[2] Qi, Wu and Zhang (2000) distinguish 5 types of shares by further sub-dividing the A and B share according to ownership restrictions.

addition to A and B shares, some Chinese companies have shares listed on foreign

stocks exchanges such as H shares listed on the Hong Kong stock exchange.

Initially the scale of the market was very small with the SHSE listing 11

companies with a total value of RMB 500 million and the SZSE listing only 5

companies with a value of RMB 270 million. However, as shown by the data

reported in Tables 1 and 2, subsequent growth was rapid, with the number of

companies listing both A and B shares rising to 530 five years after the establishment

of the exchange with 85 of these companies listing only B shares. In addition, the

number of companies listing on the Hong Kong exchange has also risen steadily to

more than 150 at present.

[Tables 1 and 2 near here]

Table 2 provides some data on the size of China's stock market since shortly

after its inception until the present. It shows strong growth in new financing, turnover

and market capitalisation. The capitalisation measure needs to be interpreted with

care, however. A hangover from state ownership of all enterprises is the continuing

high level of state ownership of listed shares which are effectively not tradable. This

is illustrated in Table 2 which distinguishes between total market value and tradable

market value with the difference between them being the shares held by government

and government-related bodies which are not traded. Clearly this is a sizable

proportion of total value – of the order of 75% of total shares by value are not tradable

– and the source of considerable market anxiety given that the government has on two

occasions attempted to begin unloading these shares with dramatic effects on share

prices leading to a rapid reversal of policy. As a measure of the size of the stock

market, the second figure is probably more appropriate than the first whereas the first

may be a better measure of the size of the listed corporate sector. Whichever measure

is used, however, it is clear that growth has been very rapid, although admittedly from a low initial base.

While the stock market has expanded rapidly since its establishment, it is still a relatively small part of the financial system. Table 3 provides some summary data and shows that the size of the stock market relative to GDP is of the order of 40-45% which compares to a value of around 150% for developed economies such as the US and the UK.[3] But even this is likely to grossly overstate the case for China – if we include only tradable shares, the ratio fell to around 15% in 2008. Thus by the standards of developing economies, the Chinese share market is very small relative to the size of its economy. In contrast, the ratio of loans to GDP is about 15 times this magnitude. If we focus on new financing, the gap is even larger. Data in Tables 2 and 3 suggest that new financing from the stock market is approximately 1.2% of GDP while loans from banks, etc. is over 100% (for 2008). All in all, the stock market, while growing rapidly, is still a relatively small part of the Chinese financial system compared to other financial institutions, primarily the banks.

[Table 3 near here]

IV The data

Since the focus of the paper is on the interrelationships between output and stock prices, we need only two series – one for stock prices and one for output. For stock prices, we used the Composite Index for the Shanghai stock exchange (the larger of the two exchanges) and for output we used GDP. Given that GDP is available only at a quarterly frequency, we restricted our analysis to quarterly data.

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[3] Comparative international data are from the World Federation of Exchanges website: www.world-exchanges.org.

The sample period used is 1992(1) to 2008(4), the start of the sample being dictated by the availability of stock price data.

The stock price data were obtained from the GTA-CSMAR data base. Data in this data base are reported on a daily basis and were averaged over the quarter to obtain quarterly observations. This was done (in contrast to, say, taking a single observation during the quarter) in order to match the GDP data best since GDP applies to each quarter as a whole. Moreover, averaging over the quarter removes very high frequency movements in stock prices which are hardly likely to respond to GDP variations (or *vice versa*). The GDP data were taken from the Wind data base for the years 1992 to 2002 and from the China State Statistical Bureau's "Financial Almanac", various years, for the period since 2002.

Neither series was seasonally adjusted. This was particularly obvious for the GDP data which has strong seasonal fluctuations. We experimented with various methods of seasonal adjustment for the GDP data, some of which were clearly unsatisfactory since there remained distinct seasonal movements in some years. We eventually undertook the main analysis with data adjusted using the X12 procedure available in EViews. Even this sophisticated method did not seem to completely remove seasonal influences so we checked our results for sensitivity to this seasonal adjustment method and briefly report the results of using an alternative method later in the paper.

Before undertaking the analysis of the relationship between our two variables, we tested them for stationarity using standard augmented Dickey-Fuller (ADF) tests. They are tests of the hypothesis that $\alpha_1 = 0$ in equation (1):

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 t + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t \qquad (1)$$

where y is the variable of interest, Δ denotes the first-difference operator, t is a time trend and ε is a random error term assumed not to be autocorrelated. Two choices need to be made before carrying out the test; first, which of the deterministic components (the intercept and the time trend) to include and, second, the lag length (the value of k). For our data these choices were not very important since the test outcome was, in general, not sensitive to either choice.

We analyse both variables in logs, as is customary, so that the first difference has the interpretation of a continuous rate of change: the rate of capital gains for stock prices and the growth rate for GDP. In Table 4 we report the ADF test statistics together with their p-values for lags 1 to 4 and various combinations of the deterministic components.[4]

[Table 4 near here]

It is clear from the results reported in Table 4 that ly is non-stationary irrespective of the nature of the deterministic components and the number of lags. The results for ls are not so clear, however. If the trend is omitted from the testing equation, ls is stationary if lags are set at 0 and 3 and when a trend is included it is stationary at lags 3 and 4. The trend terms is marginally significant so that we focus on the second line of results where the non-stationary null hypothesis cannot be rejected for lags 0 to 2 but is clearly rejected for longer lags. The SIC indicates 3 lags but tests suggest that 1 lag is sufficient to remove autocorrelation and at 1 lag ls is clearly non-stationary. We therefore conclude that, while the evidence is mixed, on balance there is evidence of non-stationarity and we proceed under the assumption that both variables are non-stationary and proceed to test for the stationarity of the first differences.

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[4] Note that we omit two deterministic cases – the one with no intercept in the levels test and the case with trend in the first-difference test.

The second part of the table suggests that first differences are stationary for both variables, although there is some doubt about the log of GDP which is non-stationary when the intercept is omitted from the ADF equation and the number of lags is chosen by the SIC. Inspection of the estimated equation, however, shows that the intercept is significant in this equation and, at SIC-determined lags, the first difference in log GDP is stationary. We conclude, therefore, that both variables are I(1).

V Results

Given our finding above that stock prices and output are both I(1), we need to determine whether they are cointegrated before we can decide on whether to model them as a VAR model in first differences or as a VECM. We proceed, therefore, to a consideration of cointegration between the two variables.

We use the Johansen test to address the cointegration question and in this case there are also numerous different possible specifications of the testing framework depending on lag length and the specification of the deterministic components of the VECM used for the test. An inspection of the data suggests that we should include an intercept in the cointegrating equation since the levels of the series are quite different. Also, a constant in the VECM seems appropriate since GDP in particular has a marked positive trend and we should allow for drift over time. Since both variables have a positive trend over the sample period, it is unlikely that a trend term will be required in the error-correction term of the VECM although we will report some results with a trend to gauge the sensitivity of the test outcomes to this assumption.

Cointegration test results for these three specifications for lag lengths from 1 to 5 are reported in Table 5.

The test outcomes are clearly dependent on the specification, both on the nature of the deterministic components and on the lag length so that a careful choice of each is required.

We begin with lag length. Tests of lag exclusion both for a VAR model in the first differences and for a VECM show that the fifth lag can be excluded but the fourth cannot. Inspection of the estimated equations reveals a strongly significant coefficient on the fourth lag of the first difference of the log of GDP in its own equation (although lags 2 and 3 are generally insignificant) which may be an indication of left-over seasonal effects in the seasonally-adjusted GDP series mentioned in section 3, a matter we will return to later in the paper. Assessment of the autocorrelation in the equations of the estimated models also shows that four lags are necessary to remove autocorrelation, with the autocorrelation being particularly strong at the fourth lag, confirming the regression results. We therefore choose four lags.

An inspection of the cointegration test results for four lags in Table 5 shows clear evidence of cointegration; this outcome does not depend on the deterministic specification or on the type of test used. We therefore conclude that the log of GDP and the log of the stock price index are cointegrated and proceed to a consideration of the implications of this relationship.

We first look at the estimated cointegrating equation. A simple OLS regression of the log of GDP (*ly*) on the log of stock prices (*ls*) produces the following:

$$ly_t = 3.2713 + 0.9511\ ls_t \qquad\qquad (2)$$
$$\quad\ (4.65) \quad\ (9.73)$$

where t-ratios appear in parentheses. Hence, a 1 percent increase in stock prices is associated with a roughly equal percentage increase in GDP. Similar magnitudes of

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the slope coefficient are obtained from more sophisticated estimates such as those obtained from the Johansen procedure. If we estimate the equation within Johansen's VECM framework with four lags and an intercept in the cointegrating regression but not in the VECM, we obtain:

$$ly_t = -1.7519 + 1.6171 \ ls_t \qquad (2a)$$
$$(0.96) \quad (6.58)$$

whereas if we also add a constant to the VECM equations, the cointegrating equation becomes:

$$ly_t = -1.3231 + 1.5905 \ ls_t \qquad (2b)$$
$$(6.53)$$

In both of these the slope coefficient is larger but of the same order of magnitude as in the simple OLS estimate. All suggest that an increase in share prices of precent is associated with an increase in real output of between 1 and 1.6 percent. Whether these are plausible or not depends on the direction in which we imagine the causation to run: a 1 percent increase in stock prices leading to a 1.6 percent increase in GDP seems very large but a 1.6:1 relationship in the opposite direction is perhaps smaller than many would expect. However, these equations cannot be used to infer causation, something to which we now turn.

To address the direction and strength of causation between our two variables, we must turn to the estimated VECM. Before we do, however, we briefly consider the concept and measurement of causation in a cointegrated system of non-stationary variables.

Testing for (Granger) causality is straightforward for pairs of stationary variables: to test for causality from x to y, simply estimate an equation explaining y in terms of lags of y and x and test for the joint significance of the lagged x variables. For pairs of non-stationary non-cointegrated variables, it is also straightforward:

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simply difference the variables until each is stationary and then apply the standard test to the differenced variables. For pairs of non-stationary cointegrated variables we can also difference and then apply the standard test to the stationary differenced series. But this ignores the error-correction term which should be added to the VAR in the differences (thus making it a VECM) and therefore carries out the test within a mis-specified model. Moreover, and more importantly, it ignores the information in the error-correction term and in particular that this terms also contains lagged x. It is possible in this context to distinguish between long-run and short-run causality. This distinction builds on the common interpretation of the VECM that the error-correction term represents deviations from long-run equilibrium while the lagged first-differenced terms capture short-run adjustment effects.

Using this distinction, testing for short-run causality is straightforward: to test for causality from x to y test the joint significance of the lagged Δx terms in the VECM equation for Δy.

Long-run causality tests are less common. We propose to use one based on the work in the unpublished papers by Canning and Pedroni (1999, 2004). Although they developed it for tests of causality in cointegrated panels, the test statistics are actually derived in a single-equation context and we follow this derivation. At present the test appears to be available for only two-variable models and we exposit it for this case.

The intuition is simple. The error-correction term in the VECM consists of deviations from the cointegrating vector which describes the long-run relationship between the two variables. Since the two variables are governed by this long-run equilibrium relationship, it must be the case that a change in one variable will be associated, in the long run, with a change in the other in order to keep the relationship

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satisfied. But this is not necessarily a causal relationship. It is possible, for example, that an exogenous change in x will be followed by a change only in y or a change only in x or, more likely, by a change in both x and y to ensure that equilibrium is re-established. We can test this very simply using the significance of the coefficients of the error-correction terms in the VECM equations: if a deviation from long-run equilibrium caused by a change in x has a significant effect on dy (that is, the error-correction term in the dy equation is significant), x causes y in the long run.

More formally, Canning and Pedroni say that x causes y in the long run if a permanent shock to the x structural error has a permanent effect on y. In terms of the earlier explanation: if a permanent shock to the x structural error is adjusted to in the long run at least partly by a change in y then x causes y in the long run. As already explained, the Canning-Pedroni test for this is based on the significance of the error-correction term in the VECM equations but requires an additional restriction which is that a shock to the y innovation has a permanent effect on y itself. In Canning and Pedroni's application, they derive this supplementary condition from the theory underlying their model.[5]

The mechanics of the test may be briefly developed as follows. Consider a two-variable VECM in x and y. Define $z = (x,y)'$ and $\Delta z = (\Delta x, \Delta y)'$ where x and y are both I(1) and cointegrated. We can write the stationary vector-moving average (VMA) form of the two-equation model for x and y as

$$\Delta z_t = F(L)\varepsilon_t \qquad\qquad (3)$$

where $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t})'$ are the structural errors and $F(L)$ is a (2x2) matrix of infinite-order polynomials in the lag operator, L, $L^n x_t = x_{t-n}$. The (i,j) element of $F(L)$ is given by:

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[5] This additional restriction appears to be ignored in many applications of the Canning-Pedroni procedure, as does the restriction to two variables; see, for example, Basu, Chakraborty and Reagle (2003), Christopoulos and Tsionas (2005), Narayan and Smyth (2008), Lee (2004) and Lee and Chang (2008).

$$F(L)_{ij} = F_{0ij} + F_{1ij}L + F_{2ij}L^2 + \ldots$$

which gives the time-path of the effects of a shock to ε_j on z_i. Note that individual elements of the $F_{\tau ij}$ sequence give the effects on Δz_i so that the sum gives the effect on z_i itself. The long run effect is just the sum of these effects from $\tau = 0$ to ∞, that is

$$F(1)_{ij} = F_{0ij} + F_{1ij} + F_{2ij} + \ldots$$

or just the accumulated effects on Δz_i. Write the matrix of $F(1)_{ij}$ elements as

$$F(1) = \begin{bmatrix} F(1)_{11} & F(1)_{12} \\ F(1)_{21} & F(1)_{22} \end{bmatrix}$$

For the question of the causality running from x to y we are interested in knowing whether F_{21} is non-zero. By the Granger representation Theorem (Engle and Granger, 1987) if x and y are I(1) and cointegrated, $F(1)$ will contain a singularity of the form $F(1)\lambda = 0$ where $\lambda = (\lambda_1, \lambda_2)'$ is the vector of coefficients of the error-correction terms in the VECM equations. The condition $F(1)\lambda = 0$ has two equations, the second of which contains the $F(1)$ element of interest, F_{21} and is:

$$\lambda_1 F(1)_{21} + \lambda_2 F(1)_{22} = 0$$

The additional condition which they derived from their underlying theory that a shock to the y innovation has a permanent effect on y itself implies that $F_{22} \neq 0$. Considering all possible combinations of signs of the λs and using the implication of the Granger Representation Theorem that not both the λs are zero, it is simple to show that $F(1)_{21}=0$ if and only if $\lambda_2 = 0$. Hence we can test the null hypothesis that $F(1)_{21} = 0$ (i.e., that x does not cause y in the long run) by testing the significance of the coefficient of the error-correction term in the equation for Δy; if we reject the null, we conclude causation from x to y.

Similarly we can test that y causes x in the long run by testing the significance of the coefficient of the error-correction term in the Δx equation provided we can impose that the restriction that the long-run effect of x on itself is non-zero.

An alternative, less restrictive, approach to the testing for long-run causality is based on the paper by Toda and Yamamoto (1995). It is applicable to variables which may or may not be stationary and which may not be cointegrated (if non-stationary). The test is based on a VAR model specified in the levels of the variables even though they may be non-stationary. If this VAR model has known order of k and the highest degree of integration of the variables is d, a VAR($k+d$) is specified, estimated by OLS and standard tests for causality are carried out, but using only the first k lags.

We now return to the estimation results. Setting lag length at 4 (on the basis of previous tests), we found that the estimated VECM equations are not sensitive to the specification of the deterministic components as long as the trend is excluded. Thus, the slope coefficient in the cointegrating regression as well as the implications for short-run and long-run causality do not depend whether there is an intercept in the cointegrating regression or in the VECM equations. Moreover, all three cases are free from autocorrelation and appear to be adequate representations of the system. In Table 6 we report the intermediate case – with an intercept in the cointegrating regression but not in the VECM equations.[6] .

[Table 6 about here]

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[6] We also experimented with a trend term in the cointegrating equation. This resulted in the trend term completely dominating the relationship between output and stock prices so that the latter was not significant implying no long-run relationship between the two, despite the outcome of the Johansen test to the contrary. Despite the lack of a relationship between the two variables, the error-correction term was clearly significant in both VECM equations implying cointegration and the anomalous result that stock prices adjust much more quickly to the gap of GDP from trend than GDP itself does. We therefore did not proceed with this case.

The implications of the VECM for short-run causality are straightforward to derive and are reported in Table 7 for 4 lags and all three combinations of deterministic components.

[Table 7 about here]

It is clear that there is no short-run causality from stock prices to output while there is considerable evidence that there is causality in the opposite direction. The evidence for the latter is strongest when there is an intercept in the cointegrating regression but not in the VECM equations, marginal when there is no intercept in either and weak when there is an intercept in both. As to the effects of varying lag length, the evidence for short-run causality from output to stock prices is stronger with fewer lags while the conclusion regarding causality in the opposite direction (no causality from stock prices to output) is not sensitive lag length. We conclude that there is no evidence of causality running from stock prices to output in the short run (irrespective of the VECM specification) but considerable evidence for causality in the opposite direction.

We turn now to examine the evidence for long-run causality. We focus on the results using the Canning and Pedroni test. The results are reported in Table 8, again for 4 lags and various combinations of deterministic components. The statistics are just the estimated coefficients and corresponding t-statistics for the coefficients of the error-correction terms in the VECM equations.

[Table 8 near here]

The results paint a very clear picture – there is strong evidence for causality from output to stock prices and no evidence at all of causality in the opposite direction. For the results reported in the table there is no sensitivity to the specification of the deterministic components of the model. There is, however, sensitivity to the

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exclusion of a trend from the cointegrating equation. If this is included, there is evidence of causality in both directions at lags of 3 and 4. There is also some sensitivity to lag length – at three lags, there is weak evidence of causality from stock prices to output for one of the deterministic combinations (an intercept in the cointegrating regression but none in the VECM). Given our earlier discussion of specification, however, these are not important qualifications to the results reported in Table 8 since the results are robust for our preferred lag length and for our preferred deterministic configuration. We conclude, therefore, that there is convincing evidence of long-run causality running from output to stock prices but none for causality in the opposite direction.

The finding of one-way long-run causality conclusion is confirmed by the application of the Toda-Yamamoto test discussed above. On the basis of a 4-lag VECM, we would specify the VAR in log levels with 5 lags so that the Toda-Yamamoto tests requires estimating a VAR in 6 lags and testing the significance of the first five to assess long-run Granger causality. The results of the application to our data are that there is no evidence of causality in either direction. However, it is quite possible that this is because the long lags substantially reduce the power of the test. This is borne out by the results obtained from shorter lags – the evidence for causality from output to stock prices grows steadily as the lag length is reduced but there is little change in the p-value for the test of causality in the opposite direction as lags are dropped.

Finally, we can, of course, obtain some informal evidence of short-run and long-run causality from simulations of the model to output and stock market shocks. We capture the results of such simulations in impulse response function (IRFs). These also provide information on the sign, magnitude and timing of the inter-

relationships between the variables. In Figure 1 we report the IRFs for our standard model – one with 4 lags and an intercept in the cointegrating equation but not in the VECM.

[Figure 1 near here]

The shocks are all unit shocks and not orthogonalised. Unit shocks were chosen so that magnitudes could be compared across shocks. Orthogonalised shocks (using the Cholesky approach) and generalised IRFs were also experimented with but were little different to the ones reported, reflecting the low cross-equation correlation of the residuals in the VECM (less than 0.01).

The IRFs bear out the conclusions we have built up so far. The two variables are clearly non-stationary in that shocks to each has a permanent effect on the variable itself, although the long-run effect is much smaller for share prices. In terms of causality, in the short run there is a small positive effect running from stock prices to output and a larger effect (initially negative) running from output to stock prices. The initial negative effect of an output shock may explain the relatively weak evidence for short-run Granger causality running in this direction despite the strong long-run evidence.

In the long run, the effects of shocks are positive in both directions but there is clearly a much larger effect running from output to stock prices than *vice versa*; a unit shock to output has a long-run effect of about 3 on stock prices while a similar shock to share prices has a negligible long-run effect on output. In terms of the earlier discussion of the estimated cointegrating equation, most of any gap that appears between stock prices and output is adjusted to by stock prices rather than output before the long-run equilibrium is re-established.

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The change in the sign of the effect of an output shock on stock prices may help reconcile apparently contradictory effects in the literature in which, as pointed out earlier in this paper, some studies show a positive and others a negative effect of output shocks on stock prices; our IRFs suggest that the effects may be negative for the first two to three years before turning substantially positive. It is possible that some of the differences arise from different time horizons, although there are doubtless other influences as well.

The overall character of the IRFs is quite robust to alterative specifications. If we vary the deterministic specification by also omitting the intercept from the cointegrating regression, the IRFs look almost identical while if we include an intercept in both the cointegrating regression and the VECM equations, the pattern of effects is much the same although adjustment to the long-run position is quicker and the magnitudes of the long-run effects are smaller. The inclusion of a trend in the cointegrating equation produces anomalous effects: "own-effects" which are still positive for output, zero for share prices and cross-effects are negative in both directions in the long run.

If we reduce lag length to three, the overall conclusions are little changed: the long-run effect of an output shock on stock prices is still positive and large while the effect of a shock in the opposite direction is small (but now negative).

We conclude that our IRFs confirm the results of earlier Granger-causality testing. Moreover, the magnitude of the effect of an output shock on share prices is positive in the long run and at least an order of magnitude larger than the effect in the opposite direction. The results are, as before, reasonably robust to alternative specifications, although not to the inclusion of a trend in the cointegrating regression.

Before rationalising our results and comparing them to the limited existing literature, we return briefly to the matter of the seasonal adjustment of the output series. It was noted above that there were often highly significant coefficients at four lags in the VECM while lags 1, 2 and 3 were generally insignificant. This suggested some residual seasonality after the X12 adjustment process. An inspection of the raw data shows very distinct seasonal fluctuations as one would expect for quarterly unadjusted GDP. There appears to have been a noticeable change in the seasonal pattern in the middle of the sample period. In the first part of the sample the amplitude of the seasonal fluctuations grow steadily as one would expect but around 2000 they decrease suddenly and markedly and then increase steadily again until the end of the sample period. The X12 procedure, not surprisingly perhaps, has trouble with this and there seems to be some residual seasonal movement in the middle of the sample after adjustment. We therefore experimented with alternative adjustment procedures and found that the use of the "Tramo/Seats" procedure available on EViews produced an adjusted series without the above problems.[7] The use of this series resulted in outcomes reassuringly similar to those reported above – (i) the two series are still clearly non-stationary, (ii)they are cointegrated, (iii)generally four lags were required in the VECM but all lags were generally significant, (iv)the evidence for short-run Granger causality is weak but less so from output to stock prices than *vice versa*, (v)there is clear evidence of long-run causation from output to stock prices but none for long-run causation in the opposite direction and (vi)the IRFs are similar in shape to those in Figure 1. We capture the thrust of these supplementary results in

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[7] The EViews manual explains that "Tramo" is an acronym for "Time series regression with ARIMA noise, missing observations and outliers" and "seats" for "signal extraction in ARIMA time series". Thank heaven for acronyms! It also claims that "it is a commonly used alternative to the Census X12 program".

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the IRFs for a VECM with four lags, and an intercept in the cointegrating equation but not in the VECM. They are presented in Figure 2.

[Figure 2 near here]

We can summarise our results very simply. First, the logs of stock prices and output are cointegrated with a significant positive long-run relationship between them. Second, there is reasonable evidence of short-run Granger causality from output to stock prices but no evidence at all of causality running from stock prices to output in the short run. Third, there is strong and robust evidence of long-run causality from output to stock prices but no causality in the opposite direction. Fourth, the long-run effect on stock prices of a shock to output is at least several times larger than the effect of a stock-price shock on output. And, finally, these results are generally robust to reasonable variations in the model (both lag length and deterministic specification) and to the method of seasonal adjustment of output.

In the light of the features of the Chinese financial system described briefly in section 3 of the paper, these results are not surprising. We saw that while the Chinese stock market has grown rapidly over the past two decades, it is still small relative to the economy as a whole and provides only minor financing for new investment while banks provide by far the bulk of financial resources for investment. Thus, it is plausible that stock price fluctuations do not affect the rest of the economy through either of the usual channels of investment and consumption. It is also consistent with the results of one of the Chinese-language papers reviewed above, Liang and Teng (2005), that shows output to be sensitive to variations in banking financial variables but not to stock market fluctuations.

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Our results are somewhat at odds with the two existing English-language papers in the area. The paper by Zhao (1999) found a strong contemporaneous effect on stock prices of output shocks, both total and unexpected output changes. However, the former had a negative effect and the latter a positive effect. The negative effect is consistent with our short-run effect although we do not distinguish between expected and unexpected shocks. Moreover, Zhao does not report any analysis of the effects in the opposite direction – not surprisingly, since only contemporaneous variables were used. Hence the results are likely to be confounded by reverse causation and it therefore difficult to compare them to ours.

The results obtained by Liu and Sinclair (2008) are, however, more clearly comparable to ours since they use a similar framework. While their conclusions regarding long-run causation are similar to ours, their short-run results are the opposite of ours – they find that short-run causation runs from stock prices to output and not *vice versa*. They do not, however, report IRFs so it can not be seen how quickly these are reversed over time as they must be if there is long-run causality running in the opposite direction. Finally, Liu and Sinclair use real GDP data with nominal stock prices which seems inconsistent and it is possible that the short-run effects we have detected are mainly on the price component in the short run but on the real component in the long run which would go some way to reconciling our results to theirs.

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VI Conclusions

In this paper we have set out to analyse the relationship between GDP and stock prices in China since the establishment of the Chinese stock market in the early 1990s. We did so using time-series techniques applied to quarterly, seasonally adjusted GDP and the Composite Index for the Shanghai Stock Exchange. Both series were used in log form.

We found both series to be non-stationary in the (log) levels but stationary in the first differences. Cointegration tests pointed clearly to the existence of a positive long-run relationship between the two variables.

In our analysis of Granger causality we distinguished between short-run and long-run causality. The evidence for short-run causality was modest and favoured the causality running from output to stock prices with no evidence for causality in the opposite direction.

Evidence for long-run causality was much more clear-cut, with output clearly causing stock prices in the long run. No evidence was found to support long-run causation running in the other direction.

Simulations of the model confirmed these findings with output shocks having an effect on stock prices which was much larger than any effect stock prices might have on output.

We argued that this was consistent with the relative immaturity of the Chinese stock market and its relatively minor role in the financial system as a whole. While it is plausible for stock prices to respond to news of changes in the economy, it is not surprising to find that changes in stock prices have at best minor repercussions on the rest of the economy.

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Table 1. Listed companies on the Chinese stock exchange

	1996	1998	2003	2008
Companies which list A shares and B shares	530	851	1287	1625
Companies which only list B shares	85	106	111	109
Companies listing H shares	25	43	93	153

Data source: for 2008: http://www.csrc.gov.cn/; for other years: "China Finance Almanac" (1996,1998,2003)

Table 2. The size of the Chinese stock market (billions of RMB)

	1993	1998	2003	2008
Financing quantum	8.47	82.58	135.72	365.67
Turnover	369.8	2255.19	3211.53	26711.26
Total market value	347.43	1950.57	4245.77	12136.64
Circulating shares market value	NA	574.56	1317.85	4521.39

Data source: http://www.csrc.gov.cn/

Table 3 Financial ratios for China

Year	1994	1996	1998	2001	2004	2008
Ratio of the value of all shares to GDP (%)	7.65	13.83	23.11	44.72	23.18	40.37
Ratio of the value of tradable shares to GDP(%)	1.98	4.03	6.81	14.86	7.31	15.04
Ratio of loans from financial institutions to GDP (%)	84.36	85.92	102.51	115.41	117.94	106.45

Data source: the value of all shares and the value of tradable shares: http://www.csrc.gov.cn/; the loans from financial institutions: http://www.pbc.gov.cn/ ; GDP: Wind data base for the years 1992 to 2002 China State Statistical Bureau's "Financial Almanac", various years, for the period since 2002.

Table 4. Tests for stationarity

Variable			Lags				
		SIC	0	1	2	3	4
ly	Intercept, no trend	1	-1.5920 (0.4811)	-1.4860 (0.5345)	-1.2797 (0.6341)	-1.8476 (0.3547)	-1.1232 (0.7016)
	Intercept and trend	0	-2.7164 (0.2336)	-2.2985 (0.4287)	-2.0838 (0.5448)	-3.0793 (0.1200)	-3.040 (0.1299)
ls	Intercept, no trend	1	-2.6385 (0.0905)	-1.6355 (0.4590)	-1.4627 (0.5461)	-3.3254 (0.0177)	-1.7089 (0.4219)
	Intercept and trend	3	-2.9503 (0.1540)	-2.5767 (0.2921)	-2.3994 (0.3765)	-4.9866 (0.0007)	-3.8416 (0.0206)
Δ*ly*	None	3	-6.8824 (0.0000)	-3.8912 (0.0002)	-2.2446 (0.0250)	-1.5227 (0.1189)	-1.4534 (0.1352)
	Intercept	0	-10.1346 (0.0000)	-6.4643 (0.0000)	-4.3612 (0.0008)	-2.5787 (0.1027)	-2.3246 (0.1676)
Δ*ls*	None	0	-8.2204 (0.0000)	-6.1124 (0.0000)	-2.1024 (0.0350)	-3.7296 (0.0003)	-4.0899 (0.0001)
	Intercept	0	-8.1501 (0.0000)	-6.0768 (0.0000)	-2.0065 (0.2834)	-3.6064 (0.0083)	-4.1625 (0.0016)

Notes: *ly* and *ls* denote the logs of GDP and the stock price index respectively; Δ denotes the first-difference operator; the figures in the "SIC" column represent the number of lags chosen on the basis of the SIC criterion; "Intercept and trend", "Intercept, no trend", "Intercept" and "None" denote the specification of the deterministic components in the ADF equation; p-values for the null hypothesis of non-stationarity are in parentheses.

Table 5. Tests for co-integration

Lags	Statistic	Deterministic Specification		
		Intercept in CE no intercept in the VAR	Intercept in CE and in the VAR	Intercept and trend in CE and intercept in the VAR
1	Trace statistic	36.6371 (0.0001)	9.1596 (0.3508)	14.3759 (0.6265)
	Max-EV	29.9114 (0.0002)	6.9593 (0.4940)	8.4693 (0.7776)
2	Trace statistic	29.6691 (0.0019)	8.8946 (0.3752)	13.2218 (0.7208)
	Max-EV	22.8117 (0.0035)	7.0616 (0.4819)	7.2475 (0.8832)
3	Trace statistic	39.3640 (0.0000)	27.2697 (0.0006)	42.1642 (0.0002)
	Max-EV	26.7391 (0.0007)	23.6776 (0.0012)	32.6145 (0.0004)
4	Trace statistic	22.9220 (0.0210)	18.6790 (0.0160)	31.2505 (0.0097)
	Max-EV	17.6191 (0.0266)	17.2630 (0.0163)	17.8338 (0.0829)
5	Trace statistic	16.2087 (0.1649)	11.7551 (0.1690)	24.9061 (0.0656)
	Max-EV	11.92037 (0.1906)	11.0488 (0.1518)	14.2602 (0.2372)

Note: "CE" denotes the cointegrating equation and "EV" denotes the eigenvalue; the null hypothesis for each test is that there are zero cointegrating vectors; p-values in parentheses

Table 6: Estimated VECM: 4 lags, intercept in the co-integrating equation

error-correction term: $ec_t = ly_t - 1.6171ls_t + 1.7519$
$$[-6.58] \qquad [0.96]$$

Variable	Δly equation	Δls equation
ec_{t-1}	-0.0084 [0.55]	0.1544 [4.14]
Δly_{t-1}	-0.0956 [0.82]	-0.4736 [1.68]
Δly_{t-2}	0.1698 [1.49]	-0.3198 [1.16]
Δly_{t-3}	0.2785 [2.41]	-0.4816 [1.73]
Δly_{t-4}	0.4982 [4.10]	-0.0739 [0.25]
Δls_{t-1}	0.0388 [0.90]	0.3825 [3.68]
Δls_{t-2}	0.0258 [0.61]	0.1220 [1.20]
Δls_{t-3}	0.0178 [0.43]	0.3529 [3.55]
Δls_{t-4}	-0.0322 [-0.72]	-0.0007 [0.01]
R-squared	0.1468	0.4311
Adjusted R-squared	0.0204	0.3468

Note: variables as defined in Table 4; absolute values of t-ratios in brackets

Table 7: Tests of short-run Granger causality

Alternative hypothesis	Deterministic specification		
	No intercept in CE or VAR	Intercept in CE no intercept in the VAR	Intercept in CE and in the VAR
ly causes *ls*	2.1987 (0.6993)	2.1929 (0.7003)	1.4349 (0.8382)
ls causes *ly*	7.1973 (0.1258)	10.0788 (0.0.0391)	5.4419 (0.2449)

Note: variables as defined in Table 4; the tests are based on a VECM with 4 lags; the statistics follow from a Wald test of the restriction that the lags of the (first-differences of the) first variable are jointly insignificant in the VECM equation for the (first-differences of the) second. They are χ^2-distributed with 4 degrees of freedom under the null of no causation. Numbers in parentheses are p-values.

Table 8: Tests for long-run Granger causality

Alternative hypothesis	Deterministic specification		
	No intercept in CE or VAR	Intercept in CE no intercept in the VAR	Intercept in CE and in the VAR
ly causes *ls*	0.1668 [3.9544]	0.1544 [4.1414]	0.1549 [4.0002]
ls causes *ly*	-0.0149 [0.8674]	-0.0084 [0.5460]	-0.0138 [0.8860]

Note: variables are as defined in Table 4; the tests are based on a VECM with 4 lags; the statistics follow from a t-test of the restriction that the coefficient of the error-correction term in the equation for the first-difference of the second variable is zero. The numbers are the estimated adjustment coefficient and the corresponding (absolute) t-ratio in brackets.

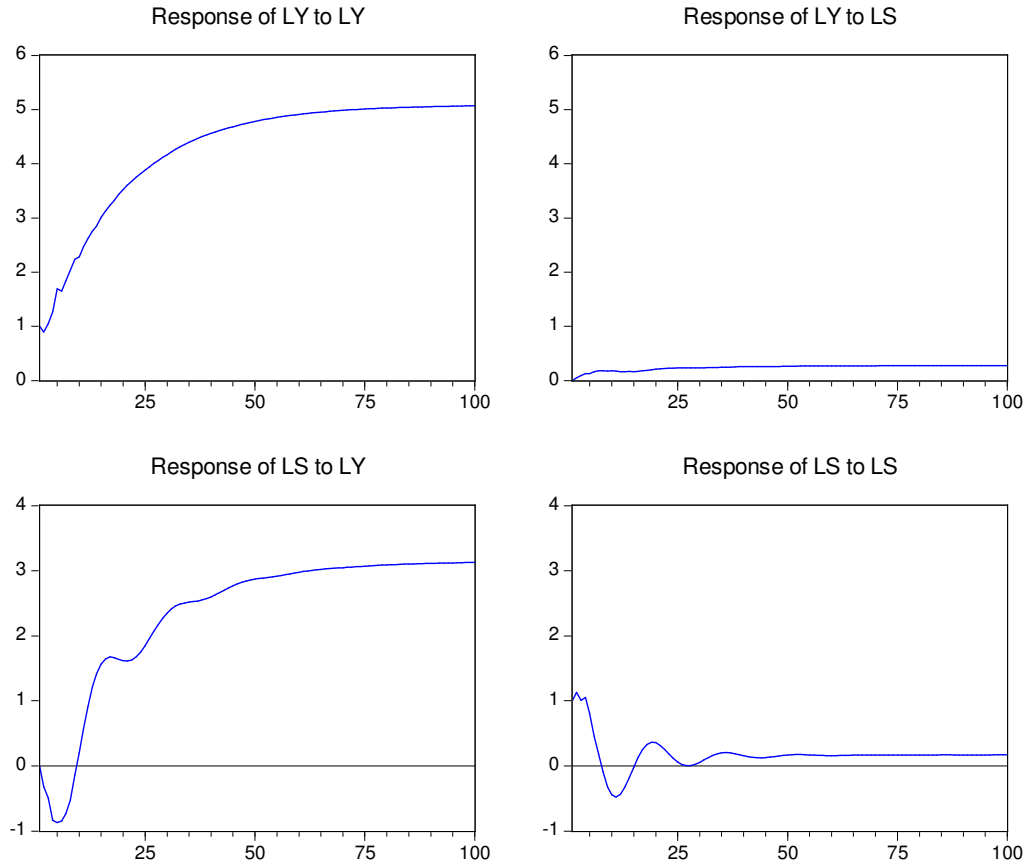Response to Nonfactorized One Unit Innovations

Figure 1 Impulse response functions from a VECM with four lags and an intercept in the cointegrating equation but not in the VECM. *ly* and *ls* are, respectively, the logs of output and share prices.

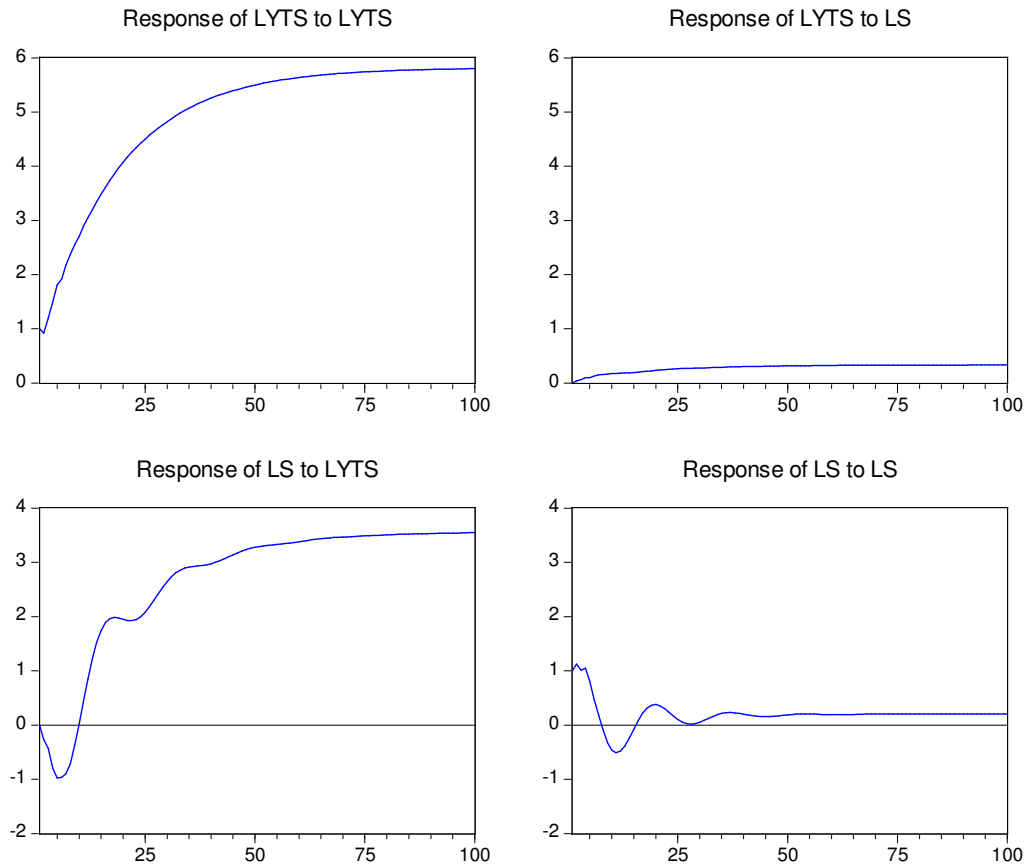Response to Nonfactorized One Unit Innovations

Figure 2. Impulse response functions from a VECM with four lags and an intercept in the cointegrating equation but not in the VECM; output seasonally adjusted using "Tramo/seats". *lyts* and *ls* are the logs of output (adjusted using Tramo/seats) and share prices, respectively.