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Banking Liquidity and Stock Market Prices in Three Countries in ASEAN

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ABSTRACT

This paper reports evidence of a banking liquidity impact on stock prices in the three Asean countries. Banking liquidity impacts suggested by Friedman is yet to be fully investigated nor verified despite several attempts. If improved liquidity of banks leads to credit expansion, which in turn leads to more positive net present value projects undertaken by firms, earnings of the latter must go up, and hence the share prices should rise. This link is worth an investigation. According to an influential of the US stock market, up to 52% of share returns are due to changes in the macro economy. Using a 3-equation structural model as well as employing corrections for cross-section dependence, we examine the link between money supply, liquidity and stock prices over 2001:4Q and 2012:2Q in three developing countries. It is found money supply changes lead to a *positive liquidity effect* and banking liquidity impacts share market prices positively. These findings are new and in support of Friedman's liquidity proposition, and also constitute evidence of a banking liquidity having a positive effect on asset prices.

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INTRODUCTION

This paper supports a six-decade old proposition of Friedman that money supply should have a positive effect on liquidity through credit expansion associated with money-to-banking-liquidity. Banking liquidity in turn, by improving firms'

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earnings, should impact on share market prices positively. Announcements of Federal Fund target rates cause large price changes in the US Treasury market (Fleming & Remolona, 1997). More than 30% of identifiable events that caused a large immediate price change in the stock market were due to monetary announcements (Fair, 2002). King, (1966) in a study, suggests up to 52% of share prices are from market-wide common factors while another study (Bernanke & Kuttner, 2005) says a 25-basis point cut in the federal funds target rates caused a one percent stock price increase. Few events are watched by market players with avid interest than monetary authority announcements. This research arises from the interest to study the role of monetary news on asset prices. Friedman's proposed positive money supply influence on inflation and a negative influence on interest rates have been empirically verified. Not so the third proposition of a money-supply-liquidity (credits) impact on asset prices. Fluctuations in share prices affect a firm's cost of capital and also its capacity to raise new capital and invest, and hence, through the wealth effect on consumption and economic growth, money supply should influence share prices, which is the subject of this paper. There is a gap in knowledge about this effect and our motivation is to find evidence to support this prediction for three countries in the Asia Pacific region.

In tracing the impact of monetary policy, the first step in the transmission channels, however, is the effect of monetary policy on the share market, through bank credit expansion from money supply changes. Policy makers are keen to understand what determines the share market's reaction to policy moves. Studies have shown a significant cyclical variation in the size of the impact of monetary policy on share prices. In those studies, the size of the response of share returns to monetary shocks has been shown to be not symmetrical - it is more than twice as large in recessionary periods and during tight credit conditions.

In the finance literature on share valuation, macroeconomic forces are assumed to have systematic influences on share prices as the so-called market-wide factors, apart from the present value of expected future cash flows. Macroeconomic influence could also be studied using the arbitrage pricing theory (APT: Ross, 1976 & Chen et al., 1986) model. In the financial economics literature, the standard aggregate demand and aggregate supply (AD/AS) framework also allows for the roles of equity markets especially in the specification of money demand (see Friedman, 1988), which linkage was labelled the monetary transmission mechanism (Mishkin, 1998). These models provide a basis for the long-run relationship and short-run dynamic interactions among macroeconomic variables on stock prices.

The specific issue discussed in this paper is the monetary transmission to share market via the bank credits to earnings of firms. This paper shows how share prices in the three ASEAN countries (Malaysia, Singapore and Thailand) are influenced by liquidity factor emanating from money supply and in our specification of macroeconomic factors. This paper links the macro-economic aggregates of money supply to banking liquidity on share prices. In the long run, fundamental factors – economic *and* firm-specific - should influence pricing, although attention in theory-building has long been on fundamental factors to the exclusion of macro factors. According to Musílek (1997), if investors want to be successful, they need to focus on price-shaping macroeconomic factors. Flannery & Protopapadakis (2002) consider macroeconomic factors as important. Another factor accounting for 10% of stock returns is the industry factor, and firm-specific factors, most commonly used in valuation models, contribute to the remaining 38% in price changes.

National income, liquidity from money supply, inflation and interest rates are important macroeconomic factors. Friedman (1969) suggested that money-supply-based liquidity has a positive influence on asset prices, which is not yet fully verified. His proposition of a *negative* money supply effect on interest rate has been confirmed by a number of studies.[1]

The main objective of this paper is to support money-to-banking-liquidity to stock price effect in each of the selected three countries. These three countries have an open capital market, and have, since 1997, deregulated their economies with large-scale financial liberalisation. Our paper differs from previous studies in that we apply the new dynamic heterogeneity panel unit root and panel co-integration test to examine the dynamic relationship between stock price, liquidity and money supply. The first step in the empirical analysis is to investigate the stochastic properties of the time series involved. Hence, we performed unit root tests on a per country basis. The power of the individual unit root tests can be severely distorted when the sample size is too small (or the span of data is too short), as in individual country data set. For these reasons, the combined data set provide reliable testing with information across countries. The Johansen co-integration test is useful to determine whether observed relationship is spurious or structural. The information is combined and to perform panel cointegration tests.

Finally, since tests verified are in a structural relationship, we estimate parameters using fully modified OLS (Ordinary Least Squares) rather than OLS to estimate the cointegrating vector for the heterogeneous cointegrated panels, which helps to correct the parameter bias induced by endogeneity and serial correlation in the regressors: that is, we avoided the errors from cross-section dependence. The long-run and short-run relationships are estimated using a vector error correction model (VECM) appropriate for

[1] The literature on the liquidity effect dates back at least to Cagan & Gandolfi (1969), Gibson (1970a; b), Leeper & Gordon (1992), Goodfriend (1997), Pagan & Robertson (1995), Christiano, Eichenbaum & Evans (1996), Hamilton (1997), Thornton (2001), Carpenter & Demiralp (2006) and Thornton (2007).

heterogeneous panels. Unlike in the past studies, monetary regime changes that may cause structural relationship to shift so introducing errors in parameter estimated, are controlled by specifying regime change controls and also controls for the Global Financial Crisis, and the Asian Financial Crisis.

This rest of the paper is organised as follows: Section 2 is a very brief discussion of the money supply theory, also its variations, by focusing the discussion on (i) liquidity and (ii) share price effect. Section 3 contains data preparation, data transformation steps, and the final test models. The findings are discussed in Section 4 before the paper concludes with relevant comments in Section 5.

MONEY SUPPLY, LIQUIDITY, SHARE PRICE AND MONETARY REGIME CHANGES

A brief review of literature is provided in this section. First, we describe the liquidity effect proposition before establishing a link between banking liquidity and share prices in the context of monetary transmission mechanism.

Money Supply Effect in the Transmission Mechanism

The effect on interest rates as a result of a change in monetary policy (Friedman's interest effect) is a proven. Many studies have examined the effect of interest rate in the monetary transmission mechanism. Such effects, following standard textbook specifications, could be represented in a money demand and money supply relationship as shown stylistically below:

$$m_t^{\,d} = \alpha_1 + \alpha_2\, r_t + \varepsilon_t^d \qquad [1.1]$$

$$m_t^{\,s} = \beta_1 + \beta_2 r_t + \varepsilon_t^s \qquad [1.2]$$

$$\therefore m_t^d = m_t^s \qquad [1.3]$$

where ^{d} indicates demand, ^{s} supply, m_t is the log of nominal money, r_t, is the nominal interest rate, while ε_t^d and ε_t^s are mutually correlated demand and supply shocks. r_t responds to shifts in money supply engineered by varying β_1 and the relation $dr_t/d\beta_1 = (\alpha_2 - \beta_2)^{-1}$ means that interest rate decreases when money supply increases, provided $\alpha_2 < 0$ and $\beta_2 < -\alpha_2$. This negative reaction to interest rate as a result of rise in money supply is termed the liquidity effect. Hence, the two-equation model as above for testing money supply and interest rates.

Share prices are expected to rise as a result of increase in money supply since a decline in interest rate would reduce the discount rate (costs of equity and debt capital) of future cash flow. This was described as a direct influence of money supply by Sprinkel (1964) for the first time. He explicitly tested a model incorporating Simple Quantity Theory (SQT) as an asset pricing model. As money expands, the portfolio of desired versus actual cash holding is thrown out of balance. Since the stock of money must be held by some agents, the prices of other assets as well as goods and services for consumption are bid up to a new equilibrium level. The channels on how the money supply influences the asset prices in the portfolio rebalancing process have newer interpretations, as for example, in Effa et al. (2013, 2011 for banking stock prices). Therefore, the relationship between money supply and stock prices is said to be positive through this adjustment mechanism on stocks.

In summary, a combination of SQT and portfolio theory is the most plausible explanation for a relationship between money supply changes and stock price changes from interest-rate-to-liquidity effect . Monetary theory is enhanced by the introduction of liquidity as it is the missing link between money and aggregate demand. Increases in liquidity can be observed during business upturns, and when money supply is eased (Quantitative Easing by the Fed from 2012 to 2015), strengthening investment, and expanding the volume of money while also enhancing financial activities. Studies by post-Keynesian economists provide new insights on money supply being endogenously rather than exogenously determined. In theoretical as well as in empirical finance, the role of liquidity has been highlighted in recent policy debates especially after the credit splurge between 1994 and 2004 that led to asset price bubbles resulting in the Global Financial Crisis (Ariff et al., 2012).

Liquidity Effect

The liquidity effect proposed by Friedman (1969) describes the first of three effects on interest rates from an unexpected change in money supply. The others are money supply effect on income as being positive and on inflation as being positive. There is a controversy (Bryant et al., 1988) as to whether money supply changes do in fact lead to negative interest rate changes as some authors conclude (Laidler, 1985). The linkage between money supply and interest rates has been recognised by policy makers on the basis of evidence of its interest rate effect more so than the unproven liquidity effect.[2] Changes in the supply of money are, therefore, a proxy for likely changes in return for money holdings.[3]

This provides a view on how the central bank uses statutory liquidity reserves to influence money supply. Among potential purchasers of assets such as institutions, dealers, and wealthy individuals with the bulk of the floating supply of corporate stock are responsive to changes in their money balances. Thus, the returns on corporate stocks will be affected, and this is called the liquidity effect. Stock prices will be responsive to movements in money supply with a negative coefficient through this channel. Despite its prominent role in conventional theories of monetary policy transmission mechanism, there has been very little empirical evidence of a

[2] The inability to find conclusive evidence has led researchers like Pagan & Robertson (1995) to suggest that this could be due to different i) definitions of money, ii) models, iii) estimation procedures and iv) sample periods.

[3] Duca (1995) adds bond funds to M2 and finds the expanded M2 provides a better explanation of the missing M2 from 1990:3 to 1992:4. As an alternative to this approach, Mehra (1997) suggested redefining the opportunity cost of M2 to include long-term bond rate to capture the increased substitution of mutual funds for bank deposits.

statistically significant or economically meaningful liquidity effect to-date.[4] It is probable that previous attempts to verify the liquidity effect have been unsuccessful because of the use of low frequency data, which necessarily mixes the effects of policy on economic variables with the effects of economic variables on policy. Hamilton (1997) sought to develop a more convincing measure of a liquidity effect by estimating the response of federal funds rate to exogenous reserve supply shocks by estimating the daily liquidity effect.

Share Prices

Where stock markets are heading and how volatile they are could well be pointed to macroeconomic as well as microeconomic factors notwithstanding the likely psychological and subjective factors. This means macroeconomic factors do have a dominant impact on share prices as suggested in the Arbitrage Pricing Theory and tested in Chen, et al. (1986).

Some of these studies examined this relationship in developed markets such as the United States, Japan and Europe (Chen, et al., 1986; Chen, 1991; Clare & Thomas, 1994; Mukherjee & Naka, 1995; Gjerde & Saettem, 1999; Flannery & Protopapadakis, 2002). Other works looked at developing markets in East Asia (Bailey & Chung, 1996; Mookerjee & Yu, 1997; Kwon & Shin, 1999; Ibrahim & Aziz, 2003). There are also studies that compare the situation in different countries (Cheung & Ng, 1998; Bilson, et al., 2001; Wongbangpo & Sharma, 2002). Studies focusing on developing markets are mostly on East Asia, to the exclusion of medium size developing economies, such as the countries included in this study.

The portfolio model of Cooper (1970) assumes that individuals could hold wealth in two forms: money and common stock. The marginal returns of stock assets determine the quantities of assets individuals will hold. A portfolio is said to be balanced when the marginal returns to holding these two assets are equal.

$$MNPS_t^M - \bar{P} = MNPS_t^S + \bar{r}_t^s \qquad [2]$$

where, the left side is the return to money asset and the right side is the return to stock asset; $\bar{P}_t$ is anticipated percentage change in general price level; $\bar{r}^*_t$ is the anticipated real pecuniary return of stocks (dividend plus change in stock prices); $MNPS_t^s$ is marginal pecuniary return to the j-th asset (the risk of j-th assets is incorporated into its pecuniary returns. $MNPS_t^M$ is implicitly a function of demand for money except for returns on alternative assets. An underlying assumption is that the positive income effect on $MNPS_t^{M,S}$ cancel each other. Thus, the

[4] A number of researchers including Bernanke & Blinder (1992) and Christiano & Eichenbaum (1991, 1992a, b) have argued that the lack of empirical evidence *is due to the Fed's attachment to interest rate targeting in one form or another* (not clear). They argued that innovations to monetary aggregates, M1, reflect shocks to money demand rather than to money supply. Consequently, the incapacity to locate the liquidity effect is due to the inability to isolate a statistically significant variable that reflect the exogenous policy actions of the Fed.

difference between $MNPS_t^M$ and $MNPS_t^{M,S}$ is primarily a function of money. In this model, money changes induce portfolio adjustments through $MNPS_t$ schedules and prices. The result is that money supply leads to stock returns.

By re-arranging this equation, the stock return can be expressed as:

$$\bar{r}_t^s = (MNPS_t^M - \bar{P}) - MNPS_t^S \quad [3]$$

Thus, Cooper's model is equivalent to the asset pricing model in finance. It would be interesting to examine the link between the liquidity effect as a result of money supply and its effect on stock prices, as proposed in this study, from Cooper's portfolio theory. Friedman's proposition could be extended as money supply having an influence on asset prices, namely share prices, in this study.

The popular finance model of equity pricing used in the industry is akin to Cooper's:

$$P_0 = \sum_{t=1}^{N} \frac{D_o\,(1+g_t)^t}{(1+i_t+r_t)^t} \quad [4]$$

where P_o is the current price of a share; D_o is the dividend at time 0; g is the constant growth rate of dividends; i_t is the risk-free rate at time t; and r_t is the equity risk premium at time t.[5] By noting the equation "D = EPS (payout)", a relationship could be shown that stock prices are correlated with EPS or some proxy such as industrial output (IPI) as representing corporate earnings, since payout ratio tends to be constant in most economies. Share prices being leading indicators of earnings is expected to lead earnings.[6] In light of current perennial financial crisis in the world, liquidity impact of money supply on stock prices has become a hot topic in policy circles to understand what ails financial systems.

The relationship between money supply and stock prices as stated by Sprinkel (1964) could also play an important role in building a model of money supply leading to asset price changes like the common stock prices.[7] Over time, interest in this topic waned until the eruption of the Global Financial Crisis, which was diagnosed as a result of *liquidity surges* that created imbalance in the financial sector and real sector (Ariff et al., 2012).

[5] By substituting EPS (payout ratio), the numerator may be replaced as (POR) . Thus, a proxy to represent EPS could be used to test if P_O is significantly affected by earnings proxy using the IPI.

[6] Chen et al. (1986) used macroeconomic variables to explain stock returns in the US stock markets. The authors showed industrial production to be positively related to the expected stock returns. Tainer (1993) is of the view that the industrial production index is procyclical. It is typically used as a proxy for real economic activity. Fama (1990) and Geske & Roll, (1983) hypothesised a similar positive relationship through the effects of industrial production on expected future cash flows. Stock & Watson (1989) have shown share prices to be a leading indicator of US economic activity as indicated by IPI and ICI in their studies covering the period 1958-1988.

[7] However, studies by Cooper (1970), Pesando (1974), Kraft & Kraft (1977), and Rozeff (1974)) have questioned this linkage between stock prices and money supply.

THE ASEAN-3 (MALAYSIA, SINGAPORE AND THAILAND) COUNTRIES

The Association of Southeast Asian Nations (ASEAN) inaugurated in 1957 with four countries now include 11 countries. The group's aim is to prevent conflict among member states by creating an integrated economic bloc through sustained modernisation. The ASEAN countries can be divided into two major categories. The inner core countries account for four percent of world trade and there are four countries in this group which have achieved greater degree of economic and financial integration among themselves and with the developed countries. The core group is generally richer and more developed than those at the periphery. There is a greater degree of financial integration among the core group with trade, economic and financial regulations similar to those in newly industrialising nations. The ASEAN group is a net exporter of merchandise and net importer of commercial services. Development of the group's financial sector has been its salient policy goal from about the mid-1980s resulting in capital market reforms.

The banks dominate the development process so the reform in this area has been the most crucial, especially because it was the banking sector that bore the brunt of the two recent financial crises, which also led to further reforms being implemented. They have been trying to diversify their heavy reliance on the banking sector in favour of direct intermediation via equity and fixed-income markets. Banks play a central role in developing countries, more than in the developed countries. This group took several measures to push for the achievement of a common ASEAN Economic Community (AEC), which was realised in 2015; for this, they adopted a scorecard to keep track of the implementation of key elements in the AEC Blueprint issued in 2007. A sanitised version of this scorecard is available in the ASEAN Secretariat website. The published version largely focuses on intergovernmental agreements, their ratification, work plans, studies, committees and other government actions.

The present study aims to find out whether there is a significant relationship between stock prices and banking liquidity arising from money supply in the three member states of Asean. We also explore how money supply affects other macroeconomic factors.

METHODOLOGY

Cointegration[8] analysis enables us to test the relationship between share prices and underlying macroeconomic variables. As mentioned in the first section, the long-run equilibrium is first examined applying a test within the VECM. Johansen cointegrating technique in this test requires the variables in the model be integrated in first order to

[8] Cointegration implies that deviations from equilibrium are stationary, with finite variance, even though the series are nonstationary and have infinite variance (Engle & Granger, 1987).

pass a test of stationarity. A further vector autoregressive VAR-like model is applied (Johansen, 1991; 1995).

Provided the residuals are I (0) or stationary, the model can be considered to be cointegrated and a valid long run relationship is assumed to exist among variables. The OLS approach, while being simple to implement, is not without problems. Parameter estimates may be biased in small samples as well as in the presence of dynamic effects; these biases are known to vary inversely with the size of the sample and the calculated R^2. When the number of regressors exceeds two there can be more than one cointegrating relationship, whence it would be difficult to give economic meaning to any finding. These problems have to be overcome so that no issues arise using the least-square estimators.

These difficulties associated with the OLS approach have led to the development of alternative procedures, the most well-known of which is proposed by Johansen (1991) using a maximum likelihood procedure to improve the OLS estimates. We adopt this procedure in our study. The existence of at most one cointegrating vector is not assumed *a prori*, but is tested for in the procedure. However, since the OLS estimator ignores the error-component structure in a model, the estimates are not efficient. Another problem with panel data modelling is when we introduce heterogeneity in order to obtain different slope coefficient for the different cross section units; the estimator bias problem emerges again for both static and dynamic models. Pesaran & Smith (1995) showed that the Fixed Effects (FE) and the Random Effects (RE) estimators will provide inconsistent results and there are two proposed solutions. One is to introduce exogenous variables into the model. If exogenous variables are added, the bias in the OLS estimator is reduced and the magnitude of the bias on the exogenous variables is biased towards zero, which is an under estimation. However, the Least Squares Dummy Variable estimator for small sample remains biased even with added exogenous variables. A second way is to use the instrumental variable methods of Anderson & Hsiao (1982).

Pool Mean Group (PMG) and Mean Group (MG)

Pesaran et al. (1999) suggest two different estimators in order to resolve the bias due to heterogeneous slopes in dynamic panels. These are: the mean group (MG) estimator and the pooled mean group (PMG) estimator. The MG method derives the long-run parameters of the panel from an average of the long-run parameters using autoregressive distributed lag ARDL models for individual countries. Using this method, they estimate separate equations for each group and examine the distribution of coefficients of these equations across groups. It provides parameter estimates by taking the means of coefficients calculated by separate equations for each group. It is one extreme of estimation because it just makes use of averaging in its estimation procedure. It does not consider any possibility of same parameters across groups.

Pesaran & Smith (1997) suggest pooled mean group estimator (PMGE) as dynamic panels for large number of observations and large number of groups. Pesaran et al. (1997; 1999) added the PMGE model. Pooled mean group estimator considers both averaging and pooling in its estimation procedure, so it is considered as an intermediate estimator. The PMGE allows variation in the intercepts, short-run dynamics and error variances across the groups, but it does not allow long-run dynamics to differ across the groups. The PMG technique is pooling the long run parameters while avoiding the inconsistency flowing from the heterogeneous short run dynamic relationships. Additionally,, the PMG relaxes the restriction on the common coefficient of short run while maintaining the assumption on the homogeneity of long run slope. The estimation of the PMG requires re-parameterisation.

Dynamic Fixed Effects Estimator (DFEE)

In addition to PMGE and MGE, Dynamic Fixed Effects Estimator (DFEE) is also used to estimate the cointegrating vector. The DFEE specification controls the country specific effects, estimated through least square dummy variable (LSDV) or generalised method of moment (GMM). The DFEE relies on pooling of cross-sections. As in the PMGE, DFEE estimator also restricts the coefficient of cointegrating vector to be equal across all panels. Adopting Pesaran et al. (1997; 1999), the PMGE model has an adjustment coefficient φ_i that is known as the error-correction term. In fact, this error-correction term φ_i indicates how much adjustment occurs in each period.

Hausman test is used to decide on the appropriate estimator between Mean Group Estimator and Pooled Mean Group Estimator. Null hypothesis test for PMGE is efficient and is consistent but MGE is inefficient against the alternative hypothesis i.e. PMGE is inefficient and inconsistent but MGE is consistent. It allows for a choice between MGE and DFEE. Therefore, we apply Hausman test on MGE, DFEE and PMGE cointegration estimates in order to decide the most efficient estimator among them. These results are supported by the Granger representation theorem (Engle & Granger, 1987) which implies that the error correction term would be significant; if significant, there is cointegration.

The long run model in (5.1)

$$SP_{it} = \mu_{it} + \beta_{1t}LQ_{it} + \beta_{2t}MS_{it} + \beta_{3t}IPI_{it} + \varepsilon_{it} \qquad [5.1]$$

will be transformed into the auto-regressive distributed lags ARDL (1,1,1) dynamic panel specification as follows:

$$SP_{it} = \mu_{it} + \lambda SP_{i,t-1} + \beta_{1t}LQ_{it-1} + \beta_{2t}MS_{it-1} + \beta_{3t}IPI_{it-1} + \eta_{it} \qquad [5.2]$$

By putting changing sign to the SP_{it}, the model in (5.2) becomes

$$\Delta SP_{it} = \mu_{it} + (\lambda\text{-}1)\ SPi_{,t-1} + \beta_{1t}LQ_{it-1} + \beta_{2t}MS_{it-1} + \beta_{3t}IPI_{it-1} + \eta_{it} \qquad [5.3]$$

From (5.2), by normalising each coefficient of the right-hand side variables by the coefficient of the SP_{t-1}, i.e. (λ_i-1) or $-(1-\lambda_i)$ since $\lambda_i<1$,

Let $\phi_i = -(1-\lambda_i)$

$$\theta_{0i} = \frac{\mu_i}{1-\lambda_i} \qquad \theta_{1i} = \frac{\beta_{10i}+\beta_{11i}}{1-\lambda_i}$$

$$\theta_{2i} = \frac{\beta_{20i}+\beta_{21i}}{1-\lambda_i} \qquad \theta_{3i} = \frac{\beta_{30i}+\beta_{31i}}{1-\lambda_i}$$

By considering the normalised long run coefficient of (5.2), the error correction re-parameterisation of (5.2) will be:

$$\Delta SP_{it} = \underbrace{\mu_{it} + \phi_i\,(SP_{i,t-1} - \theta_{0i} - \theta_{1i}LQ_{it} - \theta_{1i}MS_{it} - \theta_{1i}IPI_{it})}_{\text{Long run relations}} + \underbrace{\beta_{1t}\Delta LQ_{it} + \beta_{2t}\Delta MS_{it} + \beta_{3t}\Delta IPI_{it} + \zeta_{it}}_{\text{Short run dynamics}} \qquad (5.4)$$

The MG estimator can easily be computed from the long run parameters from the averages of the parameter values for individuals in the groups. For instance, the dynamic specification is:

$$SP_{it} = \mu_{it} + \lambda SP_{i,t-1} + \beta_{1t}LQ_{it-1} + \beta_{2t}MS_{it-1} + \beta_{3t}IPI_{it-1} + \varepsilon_{it} \qquad [5.5]$$

The long-run parameter coefficient for the equation above will be: $\theta_{0i} = \frac{\beta_{1i}}{1-\lambda_i}$

So, the entire long-run parameter will be represented as the average of long-run parameter across a group as follows:

$$\bar{\theta} = \frac{1}{N}\sum_{i=1}^{N}\theta_i \qquad \bar{\mu} = \frac{1}{N}\sum_{i=1}^{N}\mu_i$$

When the number of groups and cross sections is considerably large, the estimator for MG will be efficient even when the series is I (1). But the estimator tends to be biased when the number of time series observations is small.

The estimation of PMG and MG will be based on the model as in [5.4]. From the error correction model of (5.4), the primary interest is to see the long run coefficients (i.e. θ_{1i}, θ_{2i}, θ_{3i} and θ_{4i}). The long run coefficients provide information on the elasticity of LQ, MS and IPI factors towards the SP (stock prices) across different stock markets. Due to the uniqueness of the operation for each country, the coefficient for each factor might vary across the panels. As the coefficient provides long term equilibrium, it contains the theoretical information which is very important for each country's policy making. The error correction speed of adjustment, $\phi_{i,}$ also provides significant information to the investors: ϕ_i in equation [5.4] provides information on how long it is needed for the short run dynamics to return to long run equilibrium.

In normal situations, the short run coefficient will usually stay away from the long run equilibrium due to seasonality effect (noise), economic boom or recession. But this temporary effect as explained by the short run dynamic will eventually return to the long run equilibrium. The positive sign of the ϕ_i implies return to the long run relationships (Blackburne & Frank, 2007) from points above the regression line. The negative ign also shows the return to long run equilibrium but in opposite direction (from below). The ϕ_i is expected to be statistically significant since an insignificant coefficient of ϕ_i (i.e. $\phi_i = 0$) implies the absence of long run equilibrium. When the long run equilibrium does not exist, then is no theoretical information that could be extracted from the analysis.

Panel Unit Root Tests

Our panel dataset has time dimension of 10 years which is composed of a substantial length of time series and therefore, the existence of unit roots in variables cannot be ruled out. To test for unit root in panel data, Maddala & Wu (1999) and Choi (2001) suggest a non-parametric Fisher-type test which is based on a combination of the p-values of the test statistics for a unit root in each cross-sectional unit (the ADF-test or the PP-test).

Let p_i are U [0,1] and it is independent, and $-2\log_e p_i$ has a χ^2 distribution with 2N degree of freedom, this can be written in Equation (6.1).

$$P_\lambda = -2\sum_{i=1}^{N} log_e p_i \qquad [6.1]$$

where:

P_λ = Fisher (P_λ) panel unit root test;

N = all N cross-section;

$-2\sum_{i=1}^{N} log_e p_i$ = it has a χ^2 distribution with 2N degree of freedom.

In addition, Choi (2001) demonstrates in Equation (6.2):

$$Z = (1/\sqrt{N_{i=1}})[\sum_{i=1}^{N} \varphi_i^{-1}(p_i)] \dashrightarrow N(0,1) \qquad [6.2]$$

where:

Z: Z-statistic panel data unit root test;

N: all N cross-section in panel data;

φ_i^{-1}: the inverse of the standard normal cumulative distribution function; and

p_i : it is the P-value from the *i*th test.

Both Fisher (P_λ) Chi-square panel unit root test and Choi Z-statistics panel data unit root test have non-stationary property. The null hypothesis is:

H_0: *null hypothesis as panel data has unit root (assumes individual unit root process) and*

H_1: *panel data has no* unit root

If both Fisher (P_λ) Chi-square panel unit root test and Choi Z-statistics panel unit root test are significant, then the conclusion is to reject the null hypothesis, meaning the panel data has no unit root. If both Fisher (P_λ) Chi-square panel unit root test and Choi Z-statistics panel unit root test are not significant, it can be inferred that the null hypothesis is accepted meaning the panel data has unit root, so the series is non-stationary. Maddala & Wu and Choi tests are based on Fisher type unit root tests that are not restricted by sample sizes (Maddala & Wu, 1999).

We use two different tests to confirm our results. Maddala & Wu (1999) argued that "… other conservative tests (applicable in the case of correlated tests) based on Bonferroni bounds have also been found to be inferior to the Fisher test.". The selection of the appropriate lag length is decided using the Schwarz Bayesian Information Criterion.

Hypotheses

It is an empirical question whether principal economic indicators such as industrial production, inflation, interest rates, Treasury bill rate, liquidity, and money supply are significant explanatory factors for share returns (Litzenberger & Ramaswamy, 1982; Keim, 1985; Hardouvellis, 1987), although King (1966), and Thorbecke & Coppock (1996) have shown that share prices are affected predominantly by macroeconomic factors up to 50% and 32% respectively. In addition, if changes in economic variables are significant and consistently priced into share prices, they should be cointegrated. If there are no significant relations between macroeconomic variables and share returns, we can conclude that the stock markets of these countries do not signal changes in real activities picked up by macroeconomic factors.

It is hypothesised that money supply (MS) is endogenously determined by economic activity, mediated via the deposit-taking institutions (Effa et al., 2013). The literature on post-Keynesian theory on endogenous money is extensive.[9] Economic activity is proxied by real gross domestic product (Y), liquidity (LQ) is endogenously determined by money supply (MS) passing through the banking institutions and the share prices (SP) are endogenously affected by liquidity (LQ). Money supply (MS) is also determined by share returns (DLSP), inflation (CPI), real GDP (Y) and Treasury bill rates (TBR). Liquidity is determined by real GDP (Y), money supply (MS) and lending rate (LR).

A system of equations comprising equations for stock returns (SP), money supply (MS) and banking liquidity (LQ), is developed to be solved endogenously as follows:[10]

$$SP_{it} = f[LQ_{it}^{-}, MS_{it}^{+}, IPI_{it}^{+}] \quad [7.1]$$

$$LQ_{it} = f[MS_{it}^{+}, Y_{it}^{+}, LR_{it}^{-}] \quad [7.2]$$

$$MS_{it} = f[LQ_{it}^{+}, Y_{it}^{+}, TBR_{it}^{-}, SP_{it}^{+}, CPI_{it}^{+}, CPI(1)_{it}^{+}] \quad [7.3]$$

where SP_{it} is aggregate share price index, LQ_{it} is liquidity as proxied by reserve money, MSit is money supply as defined by M2, IPI is industrial production index, Y is real GDP, LR is lending rate, TRB is Treasury bill rate and CPI is inflation. All variables are difference in log.

Operational Model

The structural or behavioural equations can be parameterised as:

[9] Influenced greatly *by Kaldor and Moore (1982) developed* the post-Keynesian view on money, which is today the cornerstone of the PK theory of endogenous money (Rochon, 2006). The core of this theory is that causality runs from bank lending to bank deposits, instead of the traditional notion that deposits create loans.

[10] The basis of the model in this section stems from Effa et al. (2011). Not all the variables in that paper are used in this study because the focus of this study is on liquidity and stock returns; see also Dhakal et al. (1993) on causality between money and share prices observed directly.

$\ln SP_{it} = a_0 + a_1 \ln LQ_{it} + a_2 \ln MS_{it} + a_3 \ln IPI_{it} + e_{it}$ [8.1]

$\ln LQ_{it} = b_0 + b_1 \ln MS_{it} + b_2 \ln Y_{it} + b_3 LR_{it} + z_{it}$ [8.2]

$\ln MS_{it} = c_0 + c_1 \ln Y_{it} + c_2 \ln LQ_{it} + c_3 \ln SP_{it} + c_4 TBR_{it} + c_5 \ln CPI_{it} + v_{it}$ [8.3]

Separate tests for the two following hypotheses were conducted to evaluate the above models.

H_1: *MS causes Liquidity: this follows from Friedman's proposition.*

H_2: *Liquidity causes Share Prices.*

Data and Variables

Quarterly data for all variables are from the DataStream while the macroeconomic variables are compared with the International Financial Statistics (IFS) of the International Monetary Fund (IMF) for consistency. The data period: 2001:4Q – 2012:2Q. It is important to note that income is included as an explanatory variable in some equations specified above. Real gross domestic product is used as a proxy for income and since only quarterly data are available for income, the highest frequency that could be used in all regressions is quarter.

The industrial production index (IPI) is highly correlated with national income, which in turn is known to determine the earnings of firms in a modern economy.[11] Hence, we use the log change of IPI as a proxy for earnings in the equation for asset pricing: if IPI goes up, the earnings of the firms rises. Liquidity is a difficult variable to specify. There are three alternative proxies: bid-ask spread used in market studies (Amihud & Mendelson, 1986); volume of transactions (Chordia et al., 2001; Amihud, 2002); reserve money (Gordon & Leeper, 2002). Using reserve money at the central banks appears to be a right choice because if the banking system has more money in the central bank, liquidity declines, and if it keeps less reserves, liquidity rises. Hence, liquidity is inversely related to reserves. Data for money supply, M2, values are used.[12] The Treasury bill rate and the bank lending rate are the domestic 3-month Treasury-bill rate and lending rate respectively. The MSCI stock index values reported in DataStream is widely used for stock returns, P, computed as log change and adjusted as composite index for cash outflows. The log change in consumer price index is used as a proxy for inflation, INF. The bank lending rate, LR, deposit rates, TBR, and real gross domestic product, RGDP, are also obtained. All variables are seasonally adjusted where available and transformed to logarithmic form, with the exception of interest rates, which are the country-specific local 3-month Treasury bill rate, TBR and the Lending Rate, LR.

[11] A cointegration test of GDP and IPI show that these variables are cointegrated in the long-run and therefore, IPI can be used as a proxy for earnings: Chung (2013). "A Test of their Linkage between Money Supply, Liquidity and Share Prices", an unpublished PhD thesis, Universiti Putra Malaysia, Serdang, Selangor, Malaysia.

[12] The choice of monetary aggregate has been discussed earlier and its implications on the demand for money have been discussed in Pagan & Robertson (1995), and Duca (1995) on finding the liquidity effect, and for the stock market in Parhizgari (2011) on the share price effect.

The asset pricing theory discussed in Section 2 suggests a relationship between share prices and corporate dividend streams growing at g-rate of growth. The values of g and dividends depend in the long run on the earnings of the corporations, which directly depends on IPI. Although we are testing the relationship between liquidity and share prices, there is a need to control the effect of earnings changes in the system of equations. For this, we use the IPI after some initial tests using cointegration. Once the series are tested for stationarity, we run a cointegration test with income, RGDP, against IPI which confirmed these variables are cointegrated. As is evident from the test statistics, IPI is a good proxy for earnings. So, we specify this as a control in our liquidity equation for share prices. Finally, monetary regime changes and financial crises are controlled using dummies.

Descriptive statistics

Table 1 provides a summary descriptive statistics of the variables used in the regression (MG, DFE and PMG). Table 2 summarises the results of the Maddala & Wu (1999) and Choi (2001) Fisher tests. All variables are in logarithmic form except for TBR, the treasury bill rate and LR, the lending rate of banks.

The hypotheses tested are:

H_0: *each series in the panel contains a unit root, against*

H_1: *at least one of the individual series in the panel is stationary.*

Accept null if: p-value >10% (non-stationary) and

Reject null if: p-value<10% (stationary)

Table 1
Descriptive Statistics of the Panel Variables used in Tests

	LCPI	LR	LRGDP	LRIPI	LRLQ	LRM2	LSPRICE	DLSPRICE	TBR
Mean	4.66	6.04	6.73	0.00	4.64	7.03	5.39	0.02	2.14
Median	4.66	6.00	4.79	-0.01	5.55	7.93	4.85	0.03	2.28
Maximum	4.83	7.50	11.07	0.33	6.74	9.29	7.38	0.27	4.91
Minimum	4.51	5.30	4.36	-0.29	1.76	4.26	3.93	-0.42	0.21
Std. Dev.	0.08	0.67	2.92	0.14	1.86	1.90	1.14	0.10	1.15
Skewness	0.09	0.46	0.70	0.11	-0.61	-0.54	0.61	-1.04	-0.01
Kurtosis	1.84	1.91	1.51	2.40	1.52	1.52	1.69	6.64	2.23
Jarque-Bera	7.19	10.73	22.07	2.13	19.43	17.53	16.93	92.41	3.10
Probability	0.03	0.00	0.00	0.34	0.00	0.00	0.00	0.00	0.21
Sum	587.24	760.83	848.00	0.12	585.14	886.28	679.73	2.91	269.34
Sum Sq. Dev.	0.87	56.76	1063.96	2.37	430.78	450.85	163.68	1.36	165.60
Observations	126	126	126	126	126	126	126	126	126

Note: S.D. is standard deviation. LSPRICE, LRM2, LRIPI, LRGDP, LCPI, TBR, LR and LRLQ are Stock price index, Money supply, Industrial production index, Income, Inflation, Treasury bill rate, Lending rate and Reserve money respectively. All variables are in logarithmic form except for TBR and LR. DLSPRICE is the difference in stock prices, a return on stock prices.

The variables are first-differenced and computed by ratio relative to prior observation. The Jarque-Bera (JB) test indicates that all variables (except for LRIPI and TBR) are not normally distributed (JB >5.9 and *p* value of < 0.05 rejecting the null hypothesis of normality). Most of these variables are skewed (> 0, for normality should be close to 0). A quick read of the values of these variables suggest that these are as one would expect in the panel of three, namely Malaysia, Singapore and Thai economies. For example, the Treasury bill rate over the test period in these ASEAN countries is 2.14% and the lending rate is 6.04% . An expected value within known ball park values is inflation (mean of difference in log CPI) with a mean of 4.66% for the period 2001-2012. The mean of differences in LSPRICE or the share price returns is 2% over 2001-2012, with a maximum return of 27% achieved during the bull phase (2003) and a minimum of -42% during the bear phase (2008) of the market correction.

Panel Unit Root Tests Using Phillips-Perron Fisher Tests

The results indicate that, besides LRGDP and LRLQ, the null hypothesis that the series contains a unit root cannot be rejected for all variables at the one percent and 10% level of significance respectively. For example, the Fisher Chi-square statistic and Choi Z-statistic were 2.263 and 0.941 respectively for money supply (LRM2), with p-values of 0.89 and 0.83. The p-values are above the 10% level of significance, indicating the null hypothesis cannot be rejected. The variables are then tested for stationarity in first differences. The null hypothesis that the series is non-stationary when first differenced is rejected for the same variables. This means that all variables (besides LRGDP) are integrated of order one or I (1).

Since most of the variables (except for LCPI) are found to be integrated of order one (I (1)) as is also confirmed from other tests reported in the table below, the next step is to test these series to determine whether they are cointegrated. The panel cointegration test is based on Pedroni (1999, 2004) and the results are reviewed in the next section.

Pedroni Panel Cointegration Tests for Money Supply, Liquidity and Share Price

Cointegration refers to that of a set of variable that are individually integrated of order 1, some linear combinations of these variables are stationary. The vector of the slope of coefficients that renders this combination stationary is referred to as the cointegrating vector. Thus, in effect, panel cointegration techniques are intended to allow researchers to selectively pool information on common long-run relationships from across the panel while allowing the associated short-run dynamics and fixed effects to be heterogeneous across different members of the panel. Thus, the test for the null hypothesis of no cointegration is implemented as a residual-based test of the null hypothesis:

Null: H_0: γ_i for all i

Alternative: $H_{A:}$ $\gamma_i = \gamma < 1$ for all i.

Table 2
Panel Unit root tests – Fisher Phillips-Perron tests

Fisher chi-square statistic			**Choi Z-statistic**	
Variables	Level	Difference	Level	Difference
LSPRICE	5.65265	30.1579***	-0.33782	-4.25075***
	(0.4632)	(0)	(0.3677)	(0)
LRM2	2.26398	48.5210***	0.94176	-5.72605***
	(0.8939)	(0)	(0.8268)	(0)
LRLQ	16.4100*	157.704***	-1.8000	-10.6652***
	(0.0117)	(0)	(0.0359)	(0.)
LRIPI	8.54115	136.321***	-1.00672	-10.3019***
	(0.2011)	(0)	(0.1570)	(0)
TBR	2.59157	33.7573***	0.78759	-4.50771***
	(0.8581)	(0)	(0.7845)	(0)
LCPI	5.10797	57.1310***	0.18243	-6.21384***
	(0.5300)	(0)	(0.5724)	(0)
LRGDP	17.4731***	180.647***	-2.60151***	-12.1164***
	(0.008)	(0)	(0.0046)	(0)
LR	3.24330	33.4376***	0.38150	-4.49963***
	(0.7777)	(0)	(0.6486)	(0)

Note: Δ or L denotes first difference. Both variables are taken in natural logarithms. All tests take non-stationarity as null. The table shows the individual statistics and p-values with the lag length selection of one. Intercept is included in all terms with or without first differences. Probabilities of fisher type test use asymptotic $\chi 2$ distributions while other type of tests assume asymptotic normality. Numbers in parentheses are *p*-values. ***, **denote significance at the 1% and 5% level respectively. *LSPRICE rice* is stock price index, *LRLQ* is liquidity, *LRGDP* is real gross domestic product, *LRM2* is money supply, *LRIPI* is industrial production index, *LCPI* is inflation, *TBR* is Treasury bill rate and *LR* is lending rate.

Table 3 presents the result of panel cointegration test for share price, liquidity and money supply based on Pedroni panel (v, rho, pp and ADF) and group (v, rho, pp and ADF) statistics.

Lsprice: The *p*-value for Panel PP-statistics and ADF-statistics was <1% within dimension and between dimensions respectively. Thus, we can reject the null hypothesis of no integration. We can safely say that the common and individual auto regression coefficients are cointegrated. Similarly, for Lrlq: *p*-values for group-pp for between dimension was less <1% and for lrm2, *p*-values for panel v-statistic was less <5%. Thus, we can reject the null hypothesis of no cointegration. We can safely say that the common and individual auto regression coefficients are cointegrated for share price, liquidity and money supply.

Table 3
Pedroni Residual Cointegration Test Values for the Panel

Alternative hypothesis: common AR coefs. (within-dimension)						
	Lsprice		Lrlq		Lrm2	
	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.
Panel v-Statistic	0.60132	0.2738	-0.4235	0.6641	1.83068	0.0336*
Panel rho-Statistic	-0.8081	0.2095	0.45856	0.6767	1.62933	0.9484
Panel PP-Statistic	-2.1612	0.0153*	0.09237	0.5368	0.92784	0.8233
Panel ADF-Statistic	-3.0326	0.0012*	0.52307	0.6995	1.14167	0.8732
Alternative hypothesis: individual AR coefs. (between-dimension)						
	Statistic	Prob.	Statistic	Prob.	Statistic	Prob.
Group rho-Statistic	-0.0431	0.4828	-0.828	0.2038	1.9884	0.9766
Group PP-Statistic	-1.7587	0.0393*	-2.9274	0.0017*	1.01873	0.8458
Group ADF-Statistic	-3.0095	0.0013*	-0.9376	0.1742	1.44894	0.9263

FINDINGS ON ESTIMATED COEFFICIENTS

The estimates of these carefully-run tests are presented in this section. The cointegrating equations using PMG, MG and DFE on panel data of 3 countries are presented in the following section.

Cointegrating Equations Using PMG, MG and DFE on Panel Data – a comparison

Table 4a presents the results of the long-run relationship of the modelling of share price, liquidity and money supply. All of the variables appear with both the correct sign. In Table 4, the long-run coefficients of lrm2, lrlq and lripi seem to be consistent across the three estimators for LSP for stock price. There is a negative relationship with liquidity (LRLQ) and positive relationship with money supply (LRM2) and industrial production index (LIPI). The error correction term (φ_i) is negative and is less than 1 in absolute sense. φ_i is statistically significant for MG and DFE at 1% while for PMG the value is not significant at 0.10.

However, in the short run (SR), we can see the elasticity of LSPRICE as against LRLQ, LRM2 and LRIPI. The p value for ECT is 3%, 31% and 12%. For LRLQ, it is <5%, and so we accept the null hypothesis of correlation. For money supply and industrial production index the value is > than 10%, so we accept the null hypothesis of no correlation. This means that in the short run (short term), ECT for LSP (RM) and LRIPI are not significantly affecting stock prices. We can safely say that stock prices depend on the long run equilibrium in our tests of these variables (LSP(RM)), LRM2 and LRIPI).

The ECT coefficient of 0.44 reflects the period in which LSPRIC will return to equilibrium. Here, in the long run, it will take roughly 2.3 periods, or 2 quarters (referring to our data time scale in quarter) for LSPRICE to return to equilibrium if it deviates from regression line (taken as 1 / 0.44).

Table 4a
1st Equation – Share Price (LSPRICE)

	Table 4a : Cointegration Results					
	MG		DFE		PMG	
		Long Run	Parameters			
Variables	Coefficient	p-value	Coefficient	p-value	Coefficient	p-value
LRLQ	-0.79	0.51	-0.63	0.37	-2.61	0.00
LRM2	1.59	0.01	1.14	0.00	2.61	0.00
LRIPI	0.84	0.00	1.17	0.00	1.07	0.00
	Average Convergence		Parameters			
φi	-0.44	0.00	-0.29	0.00	-0.30	0.15
		Short Run	Parameters			
ΔLRLQ	0.38	0.03	0.15	0.09	0.42	0.00
ΔLRM2	-0.58	0.31	-0.63	0.39	-0.41	0.58
ΔLRIPI	0.47	0.12	0.37	0.01	0.46	0.17
Constant	-0.44	0.36	0.09	0.85	-0.34	0.27

In Table 4b, the long-run coefficients of lrm2, lrgdp and lr seem to be consistent across the three estimators for lrlq. There is a negative relationship with money supply (LRM2) and positive relationship with gdp (lrgdp) and lending rate (lr). The error correction term (φ_i) is negative and less than 1 in absolute sense. φ_i are statistically significant for MG at 0.06 while for DFE and PMG are not significant since the p-values are 0.12 and 0.11 respectively.

Table 4b
2nd Equation – Liquidity (LRLQ)

	MG		DFE		PMG	
		Long Run	Parameters			
Variables	Coefficient	p-value	Coefficient	p-value	Coefficient	p-value
LRM2	-0.29	0.64	0.93	0.21	0.44	0.08
LRGDP	2.13	0.10	0.54	0.24	0.53	0.10
LR	1.21	0.22	0.12	0.63	0.87	0.02
	Average Convergence		Parameters			
φi	-0.36	0.06	-0.08	0.12	-0.11	0.11
		Short Run	Parameters			
ΔLRM2	0.69	0.01	0.70	0.00	0.83	0.00
ΔLRGDP	0.56	0.16	0.65	0.05	0.73	0.11
ΔLR	-0.46	0.33	-0.03	0.03	-0.30	0.30
Constant	-3.29	0.00	-0.50	0.00	-0.68	0.05

"LR" means log returns of the named variable.

Table 4c
3rd Equation – Money Supply (LRM2)

	MG		DFE		PMG	
		Long Run	Parameters			
Variables	**Coefficient**	**p-value**	**Coefficient**	**p-value**	**Coefficient**	**p-value**
LRLQ	0.98	0.00	0.14	0.43	1.01	0.00
LSPRICE	-0.26	0.38	0.11	0.27	0.13	0.02
LRGDP	2.52	0.38	1.50	0.00	-1.25	0.00
TBR	-0.12	0.35	0.02	0.68	0.00	0.76
LCPI	-3.65	0.28	-0.38	0.38	1.21	0.00
	Average Convergence		Parameters			
φi	-0.26	0.02	-0.08	0.00	-0.14	0.32
		Short Run	Parameters			
ΔLRLQ	-0.07	0.72	0.15	0.01	0.03	0.92
ΔLSPRICI	0.03	0.38	0.02	0.65	0.03	0.46
ΔLRGDP	-0.10	0.78	-0.15	0.02	0.01	0.96
ΔTBR	-0.02	0.02	0.00	0.93	-0.02	0.00
ΔLCPI	-1.03	0.22	-1.39	0.01	-1.42	0.04
Constant	0.39	0.72	-0.19	0.07	0.35	0.17

In Table 4c, the long-run coefficients of LRLQ, lsprice, lrgdp, tbr and lcpi seem to be consistent across the three estimators for LRM2. There is a negative relationship with share price (LRPRCE), Treasury bill rate (TBR) and inflation (LCPI) and positive relationships with liquidity (LRLQ) and the GDP (LRGDP)). The error correction term (φ_i) is negative and is less than 1 in absolute sense. φ_i is statistically significant for MG and DFE at 2% and 8% respectively, while for PMG is not significant.

Table 4d
Hausman Test for Selection

<table>
<tr><th>MG and DFE</th><th>MG and PMG</th></tr>
<tr><td>Ho: DFE estimator is efficient and consistent but MG is not efficieint.</td><td>Ho: PMG estimator is efficient and consistent but MG is not efficieint.</td></tr>
<tr><td>Eq1 : p-value = no values</td><td>Eq1: p-value = no values</td></tr>
<tr><td>Eq2: p-value = no values</td><td>Eq2 : p-value = no values</td></tr>
<tr><td>Eq 3: p-value = 1.0 > 0</td><td>Eq 3 : p-value = 0.88 > 0</td></tr>
<tr><td>For EQ1 and 2, no coefficients in common.and no tests conducted. For Eq3 - Since Ho is not rejected, DFE estimator is efficient and consistent than MG estimator.</td><td>For EQ1 and 2, no coefficients in common and no tests conducted. For Eq3 : Since Ho is not rejected, PMG estimator is efficient and consistent than MG estimator</td></tr>
<tr><td colspan="2">Overall Decision: Both DFE and PMG estimators are found to be more efficient and consistent than MG estimator in both Hausman tests, respectively. While PMG estimator dominates the DFE estimator because it permits heterogeneity in short run coefficients. Hence PMG estimates should be relied upon, among the three estimators.</td></tr>
</table>

In Table 5, the statistics refer to final PMG regression of individual countries. We notice that, in the long run, ECT for all countries are the same. However, this cannot be said to be the same in the short run. The short run for each country is different, due to the uniqueness of one country behaviour from the others. For LSPRICE, all countries (except Thailand) have p value of < than 1%. This means that, for all countries except Thailand, we do not accept the null hypothesis of no correlation, suggesting a liquidity effect on share prices.

Table 5
PMG Final Estimation of ECT and SR (individual countries)

Lsprice												
Country	ECT	p-val	Lrlq	p-val	Lrm2	p-val	Lripi	p-val				
Malaysia	-0.10	0.01	0.20	0.26	0.93	0.09	1.01	0.03				
Singapore	-0.71	0.00	0.64	0.35	-0.52	0.56	-0.15	0.49				
Thailand	-0.07	0.31	0.42	0.22	-1.63	0.02	0.53	0.08				
Lrlq												
Country	ECT	p-val	Lrm2	p-val	Lrgdp	p-val	LR	p-val				
Malaysia	-0.09	0.04	1.05	0.00	1.44	0.01	0.02	0.01				
Singapore	-0.24	0.00	0.66	0.00	-0.11	0.39	0.88	0.06				
Thailand	0.00	0.84	0.77	0.00	0.84	0.00	-0.03	0.20				
Lrm2												
Country	ECT	p-val	Lrlq	p-val	Lsprice	p-val	Lrgdp	p-val	tbr	p-val	Lcpi	p-val
Malaysia	-0.02	0.10	0.11	0.08	0.09	0.02	-0.53	0.02	-0.03	0.15	-0.28	0.39
Singapore	0.02	0.42	0.42	0.00	0.03	0.34	0.04	0.75	-0.02	0.07	-1.34	0.02
Thailand	-0.42	0.00	-0.46	0.01	-0.04	0.32	0.54	0.00	-0.02	0.10	-2.65	0.00

For LRLQ, (except in Thailand), all countries have p-value of < than 1%. This means that, for the two with supporting evidence, we do not accept the null hypothesis of no correlation. For LRM2, all countries (except Thailand) have p-value of > than 1%. This means that for Thailand the null is accepted, while for Malaysia and Singapore, the null of no liquidity effect is rejected.

However, the D1 for LRLQ for liquidity, D11 for LRM2, D11 for LRIPI denote some very diverse patterns as can be seen in Table 5 above. These diverse results are because of the heterogeneous panel of countries. For the dependent variable LSPRICE, LRLQ has no correlation because p-value is > 10%, (so we accept the null hypothesis of no correlation) while the variables LRM2 and LRIPI are correlated with LSPRICE (so we reject null hypothesis of no correlation because p-value is < 10%, except for Singapore.

For the dependent variable LRLQ, the LRM2 is correlated with LRLQ because p-value is < 10% (so we reject the null hypothesis of no correlation) while the variables LRGDP and LR are correlated with LRLQ (we reject the null hypothesis of no correlation) because p-value is <

10%) except for Singapore for LRGDP and Thailand for LR. For the dependent variable LRM2, the following variables are correlated:

- LRLQ is correlated with LRM2 (all countries, because p-value is < 0.10)
- LSPRICE is correlated with LRM2 (except for Singapore),
- LRGDP is correlated with LRM2 (except for Singapore),
- TBR is correlated with LRM2 (except for Malaysia and Thailand), and
- LCPI is correlated with LRM2(except for Malaysia).

CONCLUSION

Using the most recently advocated estimators PMG, MG and DFE to obtain accurate results from panel regressions of data set from three countries, this paper reports evidence of money supply effects on (i) interest rates and (ii) liquidity, as well as a liquidity effect on share prices. The literature on money supply effect has been widely followed in policy circles, yet proposition (ii) and (iii) have yet to be verified. By adopting all the refinements needed to obtain robust parameter estimates described in the methodology section of this paper by using a system of equations developed for this study, this paper has offered new findings relevant to money supply and share price literature.

In conclusion, all our estimations confirmed the existence of a long run cointegration between share prices and banking liquidity, money supply and industrial production index and that they are significant. With respect to the critical assumptions of the panel analysis and with regard to the homogeneity across the panel of countries (3 in this study), empirical evidence has shown that Pool Mean Group (PMG) is the best model for the estimation of a short run and a long run relationship of share prices, liquidity and money supply.

The evidence from this paper answers our first research question that money supply causes liquidity in the short run as shown in our second equation. Liquidity causes money supply in the long run, and this impact is statistically significant. Following this, our evidence on the second research question that liquidity causes share prices is shown to be valid for the countries that we examine. In addition, while there exists a strong link between stock price and liquidity, money supply and industrial production are significantly related in the long run, but not in the short run. This is due to the money supply effect needing more time to convert credits to investments and then secure profits in the production process.

These results, as far as we know, are the first from relatively small economies, two of which (Malaysia and Thailand) are developing countries. More studies on developing countries are needed to generalise our results to developing economies. The methodology we have developed carefully to secure econometric sophistication should be followed if the results are to be reliable given the problems of stationarity, serial correlation, country heterogeneity, regime changes, crisis-effects, all of which could introduce errors in the estimates.

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