PRICE MOVEMENTS AND PRICE DISCOVERY IN THE MUNICIPAL BOND INDEX AND THE INDEX FUTURES MARKETS

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INTRODUCTION

Municipal bonds are long-term debt instruments issued by states and municipalities. The most important advantage of municipal bonds to the investor is the exemption of interest income from federal taxation (and sometimes from State taxation). The estimated total market value for municipal bonds is as high as $12 trillion with 1.5 million issues sold by 50,000 state and municipal entities.[1]

The Chicago Board of Trade introduced the Municipal Bond Index (MBI) futures contract on June 11, 1985. It has become increasingly popular among municipal bond issuers, investors, dealers, and

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[1]See, *Business Week,* September 6, 1993, p. 45.

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arbitrageurs.[2] Surprisingly, very little research has been done on this important market although other futures markets have been investigated extensively.

This article, using the cointegration methodology developed by Engle and Granger (1987), examines the temporal price relationship between the Municipal Bond Index (MBI) and the futures contract. Particularly, the price discovery role of the index and index futures markets is investigated. If the cash and futures markets are efficient and frictionless, arbitrage (trading between markets, or basis trading) activities will maintain correct pricing relationships. The price movements should be simultaneous, otherwise profits can be extracted from the markets. However, a lead–lag price movement relationship is possible due to differences in market structure, transaction costs, market liquidity, and short sale constraints in the cash market and other regulatory and technical constraints. (See, e.g., Stoll and Whaley (1990) and Wahab and Lashgari (1993) for more detailed discussions.)

This study is interesting for several reasons. First, many studies have examined the price relationship between cash and futures markets for short-term interest rate futures, i.e., Fung and Leung (1993), and MacDonald and Hein (1989). However, long-term interest rate futures have been largely neglected in the literature. This might be due to difficulties in defining precisely what is the underlying security of the Treasury Bond futures because of various delivery options available to traders with a short position.[3] The MBI futures contract, however, is settled by the cash method. Thus, it provides an unique opportunity to study the long-term interest rate futures market without the complication of an undefined underlying security as in the case of the Treasury Bond futures contract. Second, the municipal bond market is itself interesting for investigation. Municipal securities differ from Treasury securities in many important aspects, including some unique risks, such as a changing default pattern in different economic cycles and the possibility of new tax legislation.

This study finds that the MBI and the index futures prices are nonstationary. The cointegration analysis reveals that there exists a long-run equilibrium relationship between these two nonstationary price series. The error correction model (ECM) representation of the cointegrated price series suggests that there is a two-way feedback relationship and

[2]For example, on February 18, 1994, the Municipal Bond Index futures trading volume and open interests were \$1.3 trillion and \$2.7 trillion, respectively.

[3]These delivery options are usually referred to as quality, timing, and wild card options.

that the feedback from futures to cash is the stronger of the two feedbacks. Sub-sample results indicate that the information discovery role of the futures market has strengthened during the sample period. These findings are of importance. Investors can potentially improve hedging performance by incorporating the long-run cointegrating relationship and short-term dynamics between the MBI and the index futures prices [see, Kroner and Sultan (1993)].

The next section briefly reviews the relevant literature. The third section describes the data and some summary statistics, while the fourth studies the order of integration of the MBI and the index futures price series. In the fifth section, the cointegration analysis methodology of Engle and Granger (1987) is used to examine the long-run equilibrium relationship of the two price series and in the sixth an error correction model (ECM) is used to study the short-run dynamics of the two price series.

LITERATURE REVIEW

The price discovery function of futures markets has long been recognized in theoretical literature [Garbade and Silber (1983); Grossman (1977); Working (1970)].[4] Recently, many studies have analyzed the price discovery function of futures prices in various futures markets. However, the price discovery function of futures markets is not very conclusive. The majority of these studies do support the notion that the futures market leads the cash market in incorporating new information. For example, Kawaller, Koch, and Koch (1987) investigated the intraday price relationship between the S&P 500 index and the index futures. Their results suggest that changes in the index futures lead changes in the index by 20 to 45 minutes and that the feedback from the cash market to the futures market is rather weak. Stoll and Whaley (1990) and Chan (1992) studied both S&P 500 index and Major Market Index (MMI) cash and futures markets and also found the predominant futures-to-cash lead. Conversely, for instance, Fung and Leung (1993) found that there is a two-way feedback relationship between the cash and futures markets for Eurodollar deposits. Wahab and Lashgari (1993) studied the cash and futures markets for both the S&P 500 and FT-SE 100 indexes. They found that although feedback exists between the cash and futures markets, the spot-to-future lead appears stronger than the other direction.

[4]Another function of futures markets is to provide a mechanism through which hedgers transfer price risk to speculators.

DATA AND SUMMARY STATISTICS

The Municipal Bond Index futures contracts are based on the *Bond Buyer* Municipal Bond Index. The index includes 40 actively traded, high-quality municipal bonds. The constituent bond prices are obtained by excluding the highest and lowest quotes provided by five bond dealers and averaging the three remaining quotes. The MBI futures contracts are settled by the cash method and they mature on a March, June, September, and December cycle. Daily spot and futures index prices from June 11, 1985, through April 28, 1993, are from the Chicago Board of Trade. There are a total of 1983 daily observations. The futures price series are constructed by using the contract closest to expiration.

Table I presents some summary statistics of the MBI (P_t), MBI futures (F_t), and the differenced series ΔP_t and ΔF_t for the whole sample period and two equally divided sub-samples. The mean, standard deviation (STD), minimum and maximum values as well as autocorrelation structure are reported for each level and differenced series. ρ_i, $i = 1, 2, \ldots, 6$, is the ith order autocorrelation coefficient. Panel A presents the summary statistics for the full sample. The mean value of spot price P_t is higher than that of futures price, F_t, which suggests that it is more expensive, on average, to establish a cash position than a futures position. This is not surprising because a cash position entitles one to collect coupon interest income. When coupon interest rate is higher than the short-term interest rate, the familiar cost-of-carry model predicts that P_t should be higher than F_t.[5] The autocorrelation coefficients for both P_t and F_t are very high and close to 1.0 and they die down very slowly. However, the autocorrelation coefficients for the differenced series, ΔP_t and ΔF_t, are much lower. This suggests that there may be a unit root in the level series. The summary statistics for the two sub-samples are qualitatively similar (see, Panels B and C).

INTEGRATION TEST

A nonstationary time series, X_t, is said to be integrated of order d, or $I(d)$, if it become stationary after differencing d times. Many studies suggest that cash and futures price series are nonstationary series [e.g., Baillie and Bollerslev (1989); Fung and Leung (1993); Wahab and Lashgari (1993)]. The order of integration of a nonstationary time series

[5]The yield curve for tax-exempted municipal bonds are always upward-sloping during the period tested. Salomon Brothers' *Analytical Record of Yields and Yield Spreads* (1991) provides monthly prime grade municipal yields for various maturities from 1954 through 1990.

TABLE I
Summary Statistics for Municipal Bond Index and Index Futures Series

Series	*Mean*	STD	*Min*	*Max*	*Autocorrelation Coefficient* ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6
Panel A: Full Sample, 11/06/1985–28/04/1993, $n = 1983$										
Index (P_t)	92.916	4.371	76.313	103.250	0.995	0.989	0.982	0.975	0.967	0.960
Futures (F_t)	92.278	4.618	74.219	102.750	0.992	0.983	0.975	0.966	0.957	0.949
Panel B: First Sub-sample, 11/06/1985–24/05/1989, $n = 992$										
Index (P_t)	91.743	5.117	76.313	103.000	0.996	0.989	0.982	0.975	0.967	0.960
Futures (F_t)	90.886	5.425	74.219	102.750	0.992	0.984	0.975	0.966	0.957	0.949
Panel C: Second Sub-sample, 25/05/1989–28/04/1993, $n = 991$										
Index (P_t)	94.090	3.047	87.656	103.250	0.995	0.989	0.981	0.972	0.964	0.954
Futures (F_t)	93.672	3.059	86.188	102.656	0.990	0.979	0.967	0.956	0.945	0.934
Panel D: Differenced Series, Full Sample										
Index (ΔP_t)	0.0061	0.360	−4.063	2.656	0.264	0.086	0.083	0.015	−0.056	0.013
Futures (ΔF_t)	0.0067	0.531	−3.000	3.000	0.087	−0.004	0.002	−0.020	−0.037	−0.022

can be inferred by testing for unit roots. The Dickey–Fuller (DF) and augmented Dickey–Fuller (ADF) tests are used in this study. To conduct the DF test, the following ordinary least square (OLS) regression is first performed:

$$\text{DF: } \Delta X_t = \alpha_o + \alpha t + \beta X_{t-1} + \epsilon_t \tag{1}$$

where $\Delta X_t = X_t - X_{t-1}$ and t is the time trend. ϵ_t is assumed to be an i.i.d. error term. If ϵ_t is serially correlated, the ADF test which includes lagged differences, is considered more appropriate. The ADF test is based on an OLS regression of the following form:[6]

$$\text{ADF: } \Delta X_t = \alpha_o + \alpha t + \beta X_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta X_{t-i} + \eta_t \tag{2}$$

The number of lags, k, should be chosen such that η_t is an i.i.d. error term. Both the DF and ADF tests center on the β coefficient of the lagged level term, X_{t-1}. Under the null hypothesis that X_t is a unit root process, β equals zero.

The spot and futures price series, P_t and F_t, are tested for unit roots. The regression t-statistics are reported in Table II. For the ADF

TABLE II

Test for Unit Root on Municipal Bond Index and Index Futures Series

Series	*DF Test*	*ADF Test* ($k = 4$)	*ADF Test* (k = 12)
Panel A: Full Sample, 11/06/1985–28/04/1993			
Index (P_t)	−1.720	−2.533	−2.833
Index futures (F_t)	−2.590	−2.799	−2.968
Panel B: First Sub-sample, 11/06/1985–24/05/1989			
Index (P_t)	−1.524	−2.003	−2.328
Index futures (F_t)	−1.989	−2.167	−2.468
Panel C: Second Sub-sample, 25/05/1989–28/04/1993			
Index (P_t)	−1.758	−2.761	−2.775
Index futures (F_t)	−2.695	−2.684	−2.735

Note: 1. The critical values at the 5% and 1% levels are −3.42 and −3.98, respectively. See Fuller (1976, p. 373).
2. The DF and ADF tests are based on the following OLS regressions:

$$\text{DF: } \Delta X_t = \alpha_o + \alpha t + \beta X_{t-1} + \varepsilon_t \tag{1}$$

$$\text{ADF: } \Delta X_t = \alpha_o + \alpha t + \beta X_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta X_{t-i} + \eta_t \tag{2}$$

where $X_t = P_t$ or F_t.

[6] A time trend term is included in eqs. (1) and (2). However, the DF and ADF tests without the trend term are also conducted and the results are qualitatively similar.

tests, a number of different lags, k, up to 12, are tried. The results are qualitatively the same.[7] Only the two cases with $k = 4$ and $k = 12$ are reported. It can be seen that both the DF and ADF tests suggest that the unit root hypothesis cannot be rejected for the full and two sub-samples. Hence, P_t and F_t are integrated at least of order one, possibly, of higher order.

It is necessary to ensure that the two series are integrated of the same order to apply the cointegration test. It is well known that if a level series possesses two unit-roots, the differenced series must possess one unit-root. Therefore, the same DF and ADF tests are applied to the two differenced series, ΔP_t and ΔF_t. The results are reported in Table III. It is found that the unit root nonstationary hypothesis for the two differenced series can be rejected decisively for the full sample and two sub-samples. The combined results in Tables II and III suggest that both P_t and F_t are $I(1)$ series.

COINTEGRATION TEST

The integration tests above suggest that the spot and futures index prices are nonstationary $I(1)$ series. Market efficiency implies that F_t and P_t

TABLE III

Test for Unit Root on Differenced Series

Series	*DF Test*	*ADF Test* ($k = 4$)	*ADF Test* ($k = 12$)
Panel A: Full Sample, 11/06/1985–28/04/1993			
Index (ΔP_t)	−33.93	−19.26	−11.22
Index futures (ΔF_t)	−40.77	−20.54	−11.67
Panel B: First Sub-sample, 11/06/1986–24/05/1989			
Index (ΔP_t)	−24.40	−14.20	−7.65
Index futures (ΔF_t)	−28.89	−14.76	−7.86
Panel C: Second Sub-sample, 25/05/1989–28/04/1993			
Index (ΔP_t)	−22.00	−10.98	−8.52
Index futures (ΔF_t)	−28.58	−13.66	−9.20

Note: 1. The critical values at the 5% and 1% levels are −2.87 and −3.44, respectively. See Fuller (1976, p. 373).
2. The DF and ADF tests are based on the following OLS regressions:

$$\text{DF: } \Delta S_t = \alpha_o + \alpha t + \beta S_{t-1} + \varepsilon_t \tag{1'}$$

$$\text{ADF: } \Delta S_t = \alpha_o + \alpha t + \beta S_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta S_{t-i} + u_t \tag{2'}$$

where $S_t = \Delta P_t$ or ΔF_t.

[7]When k is unknown, Said and Dickey (1984) suggest that it should increases with the sample size, T, at the rate of $T^{1/3}$.

should not deviate apart without bound and that economic forces should prohibit persistent long-run deviations from the equilibrium between cash and futures prices of a same asset [Garbade and Silber (1983); Granger (1986)]. In this section, the cointegration analysis methodology is used to test whether there exists an equilibrium relationship between these two series.

Two $I(1)$ series, X_t and Y_t, are said to be cointegrated if there exists a linear combination of the two series for some constant, a, such that $X_t - (c + aY_t) = Z_t$ is a stationary process, i.e., $I(0)$. Z_t is called the equilibrium error. Test for cointegration can be carried out by checking for stationarity of the residuals from the following OLS cointegration regression:

$$X_t = c + aY_t + Z_t \tag{3}$$

where c is a constant; a is the estimator of the cointegration parameter; and Z_t is the error term. If X_t and Y_t are cointegrated, Z_t is a stationary process.[8] Under the null hypothesis of non-cointegration, Z_t should possess a unit root. Engle and Granger (1987) recommend the ADF procedure to test for unit root in Z_t series, which requires the estimation of the following OLS model:

$$\Delta Z_t = a_o + b_o Z_{t-1} + \sum_{i=1}^{l} b_i \Delta Z_{t-i} + e_t \tag{4}$$

where a_o, b_i, $i = 0, 1, 2, \ldots, l$, are regression coefficients and e_t is the error term. Following the model-building strategy adopted in Engle and Granger (1987, p. 271–274), the regression eq. (4) is first fitted with 12 lagged changes of Z_t, i.e., 1 is set to 12. The test results are reported in Table IV. The non-cointegration null hypothesis can be rejected at the 1% level. It is found also that the regression coefficients b_i for $i = 5, 6, \ldots, 12$, are insignificant.[9] Hence, a modified regression eq. (4) with $1 = 4$ is fitted again. The t-statistics for the modified model are higher (in terms of absolute value) than those with 12 lagged changes. The null hypothesis of non-cointegration can be rejected at the 1% level. It is interesting to note that the ADF tests with four lagged changes of residuals are more powerful than those with 12 lags. This is because the

[8]A constant c is included to allow for the possibility that Z_t has a non-zero mean. In the empirical application, $X_t = P_t$, $Y_t = F_t$ and the equilibrium error is denoted as z_t^P, when P_t is the LHS variable. $X_t = F_t$, $Y_t = P_t$ and the equilibrium error is denoted as Z_t^F when the LHS is F_t.

[9]An alternative model-building strategy used in some empirical studies is to use Akaike's final prediction error (FPE) criterion or Shwartz information criterion to determine the appropriate number of lags. This study finds that the optimal number of lags is 4 in all three cases.

introduction of extraneous insignificant parameters decreases the power of the tests. The results in Table IV show that there exists a long-run equilibrium relationship between P_t and F_t, even though each of the two series are nonstationary individually. This is also true for the two sub-samples.

ERROR CORRECTION REPRESENTATION

Series Transformation

The evidence from the integration tests section suggests that both the index spot and futures prices are nonstationary $I(1)$ series. The following transformation is used to achieve stationarity:

$$p_t = \ln P_t - \ln P_{t-1}$$
$$f_t = \ln F_t - \ln F_{t-1} \tag{5}$$

where p_t and f_t represent the (continuously compounded) daily rates of return of the index spot and futures series.

TABLE IV

Cointegration Tests for Municipal Bond Index and Index Futures Series

Series	c	a	R^2	t-*Statistic* *for* b_o (1 = 4)	t-*Statistic* *for* b_o (1 = 12)
	Panel A: Full Sample, 11/06/1985–28/04/1993				
Index (P_t)	6.656	0.935	0.975	−7.625[a]	−6.653[a]
Index futures (F_t)	−4.667	1.043	0.975	−7.789[a]	−6.856[a]
	Panel B: First Sub-sample, 11/06/1986–24/05/1989				
Index (P_t)	7.039	0.932	0.976	−5.683[a]	−5.270[a]
Index futures (F_t)	−5.211	1.047	0.976	−5.777[a]	−5.405[a]
	Panel C: Second Sub-sample, 25/05/1989–28/04/1993				
Index (P_t)	2.118	0.982	0.971	−5.674[a]	−4.232[a]
Index futures (F_t)	0.594	0.989	0.971	−5.864[a]	−4.418[a]

Note: The cointegration analysis is based on the following regression equations:

$$X_t = c + aY_{t-1} + Z_t \tag{3}$$

$$\Delta Z_t = a_o + b_o Z_{t-1} + \sum_{i=1}^{l} b_i \Delta Z_{t-i} + e_t \tag{4}$$

When P_t is the LHS variable in (3), $X_t = P_t$, $Y_t = F_t$ and the equilibrium error is denoted as Z_t^P subsequently; when F_t is LHS variable in (3), $X_t = F_t$, $Y_t = P_t$ and the equilibrium error will be denoted as Z_t^F.

[a]Indicates the test statistic is significant at the 1% level. The critical value is from Engle and Granger (1987).

The observed price series may not represent the true price levels due to bid–ask spreads or infrequent trading. There may also exist a nonsynchronous problem in the underlying municipal bonds of the index. Following Stoll and Whaley (1990) and Wahab and Lashgai (1993), time series models are fitted first to purge the noises in p_t and f_t. An ARMA model with various parameterizations are tried. It is found that the following parsimonious MA(1) model fits p_t and f_t well:

$$x_t = m^x + u_t^x + \theta^x u_{t-1}^x, \qquad x = p, f \tag{6}$$

where m^x, θ^x, and u_t^x are the mean, the moving average parameter, and the process innovation at time t, respectively. Table V presents the serial correlation structure before and after purging the noises in the transformed series, p_t and f_t.[10] It can be seen that the serial correlation coefficient is reduced substantively, especially for initial lags,

TABLE V

Sample Serial Correlation Coefficients of the Transformed MBI Spot and Futures Prices and the Innovations Generated by Estimated MA(1) Processes

	Transformed Series		*Series Innovations*	
Lag	*Index p_t*	*Futures f_t*	*Index u_t^p*	*Futures u_t^f*
1	0.263	0.088	0.013	−0.000
2	0.082	0.002	0.065	0.002
3	0.080	0.004	0.063	0.005
4	0.009	−0.022	0.011	−0.020
5	−0.063	−0.038	−0.071	−0.034
6	0.009	−0.028	0.018	−0.028
7	0.035	0.034	0.033	0.038
8	−0.004	−0.016	−0.007	−0.017
9	−0.011	−0.026	−0.019	−0.026
10	0.031	0.014	0.041	0.015
11	0.005	0.013	−0.023	0.007
12	0.068	0.063	0.086	0.062

Estimated MA(1) processes

$$p_t = 0.0000649 + u_t^p + \underset{(11.55^a)}{0.251} u_{t-1}^p \qquad R^2 = 0.065$$

$$f_t = 0.0000715 + u_t^f + \underset{(3.95^a)}{0.088} u_{t-1}^f \qquad R^2 = 0.008$$

Note: The number in parentheses are the t-statistic.
[a]Indicates that the statistic is significant at the 1% level.

[10]Sub-sample results are very similar. They are not reported to conserve space.

after purging the noises in the observed series.[11] The estimated moving average parameters are statistically significant. The t-statistic of the moving average parameter for series p_t is much higher than that for series, f_t. The regression R^2 for the index spot series is also much higher than that for the index futures series. These results are not surprising because one would have expected the futures prices to be less contaminated by trading noises. A non-synchronous trading problem does not exist and an infrequent trading problem is much less serious in index futures since it is traded as a single asset.

Error Correction Model Estimation

The evidence from the cointegration test section indicates that there exists a long-run equilibrium relationship between the index spot and futures series. Engle and Granger (1987) have proved that a cointegrated series must have a corresponding error correction model (ECM) representation. In this section, the following ECM is used to investigate the short-run dynamics of the two price series.

$$\text{ECM: } x_t' = \alpha_o + \beta_o Z_{t-1} + \sum_{i=1}^{m} \beta_i x_{t-i}' + \sum_{j=1}^{n} \gamma_j y_{t-j}' + v_t \qquad (7)$$

where x_t' and y_t' represents the innovations of the transformed spot and futures series in eq. (6), i.e., $x_t' = u_t^x$, $y_t' = u_t^y$; Z_{t-1} is the previous period's equilibrium error, i.e., the regression residuals from eq. (3); β_i and γ_j are regression coefficients; v_t is the error term. The appeal of the ECM representation is that it combines the short-run dynamics (responses from lagged innovations, x_{t-i}' and y_{t-j}') with a long-run equilibrium relationship ($X_t = c + aY_t$). This equation enables one to investigate whether the futures price is able to forecast the cash price, or vice versa; and whether the futures market or the cash market performs most of price discovery rule.

The results of OLS estimation of the error correction models are reported in Table VI. Four error correction regressions are estimated for the whole and each of the two sub-sample periods: (a) index spot innovations are regressed on last period's spot (own-market) equilibrium error and lagged innovations in spot and futures series; (b) index spot innovations are regressed on last period's futures (cross-market)

[11]One should be concerned more with correlation coefficients at the initial lags, especially at lag one since this study deals with daily, not minute-to-minute, data. Some correlation coefficients are still marginally significant at distant lags. One can not assume that they result from bid–ask spreads or infrequent trading.

TABLE VI

OLS Estimation Results of Error Correction Models

Panel A: Full Sample, 11/06/1985–28/04/1993

$$p'_t = \underset{(0.477)}{0.000} - \underset{(-2.344^{a})}{0.00031Z^{P}_{t-1}} - \underset{(-7.923^{b})}{0.278p'_{t-1}} - \underset{(-3.699^{b})}{0.125p'_{t-5}} + \underset{(9.651^{b})}{0.226f'_{t-1}} + \underset{(4.023^{b})}{0.060f'_{t-2}} + \underset{(3.993^{b})}{0.058f'_{t-3}} + \underset{(2.260^{a})}{0.050f'_{t-5}} + \underset{(3.600^{b})}{0.051f'_{t-12}}$$

S.E.E. $= 0.0037$ DW $= 2.009$ $F(8, 1961) = 21.38^{b}$ $R^2 = 0.080$

$$p'_t = \underset{(0.477)}{0.000} + \underset{(-2.022^{a})}{0.00025Z^{F}_{t-1}} - \underset{(-7.965^{b})}{0.279p'_{t-1}} - \underset{(-3.733^{b})}{0.126p'_{t-5}} + \underset{(9.724^{b})}{0.228f'_{t-1}} + \underset{(4.062^{b})}{0.061f'_{t-2}} + \underset{(4.008^{b})}{0.058f'_{t-3}} + \underset{(2.284^{a})}{0.051f'_{t-5}} + \underset{(3.586^{b})}{0.051f'_{t-12}}$$

S.E.E. $= 0.0037$ DW $= 2.010$ $F(8, 1961) = 21.19^{b}$ $R^2 = 0.080$

$$f'_t = \underset{(0.696)}{0.000} - \underset{(-2.854^{b})}{0.00052Z^{F}_{t-1}} - \underset{(-2.872^{b})}{0.098p'_{t-5}} - \underset{(-2.328^{a})}{0.124p'_{t-11}} + \underset{(2.029^{a})}{0.071f'_{t-11}}$$

S.E.E. $= 0.0058$ DW $= 1.961$ $F(4, 1966) = 5.59^{b}$ $R^2 = 0.011$

$$f'_t = \underset{(0.690)}{0.000} + \underset{(-2.480^{a})}{0.00048Z^{P}_{t-1}} - \underset{(-2.906^{b})}{0.099p'_{t-5}} - \underset{(-2.332^{a})}{0.124p'_{t-11}} + \underset{(2.011^{a})}{0.070f'_{t-11}}$$

S.E.E. $= 0.0059$ DW $= 1.966$ $F(4, 1966) = 5.08^{b}$ $R^2 = 0.010$

Panel B: First Sub-sample, 11/06/1985–28/04/1993

$$p'_t = \underset{(0.276)}{0.000} - \underset{(-2.764^{b})}{0.00057Z^{P}_{t-1}} - \underset{(-5.900^{b})}{0.303p'_{t-1}} - \underset{(-3.332^{b})}{0.102p'_{t-5}} + \underset{(6.892^{b})}{0.250f'_{t-1}} + \underset{(2.213^{a})}{0.049f'_{t-2}}$$

S.E.E. $= 0.0048$ DW $= 2.013$ $F(6, 972) = 16.62^{b}$ $R^2 = 0.078$

$$p'_t = \underset{(0.201)}{0.000} + \underset{(2.563^{a})}{0.00051Z^{F}_{t-1}} - \underset{(-5.881^{b})}{0.301p'_{t-1}} - \underset{(-3.467^{b})}{0.106p'_{t-5}} + \underset{(6.871^{b})}{0.249f'_{t-1}} + \underset{(2.176^{a})}{0.049f'_{t-2}} + \underset{(3.176^{b})}{0.068f'_{t-12}}$$

S.E.E. $= 0.0048$ DW $= 2.006$ $F(6, 972) = 15.47^{b}$ $R^2 = 0.087$

$$f'_t = \underset{(0.408)}{0.000} - \underset{(-1.262)}{0.00035Z^{F}_{t-1}} - \underset{(-3.422^{b})}{0.156p'_{t-5}} + \underset{(\ 2.677^{b})}{0.085p'_{t-12}}$$

S.E.E. $= 0.007$ DW $= 1.981$ $F(4, 975) = 6.55^{b}$ $R^2 = 0.019$

$$f'_t = \underset{(0.458)}{0.000} + \underset{(\ 1.041)}{0.00031Z^{P}_{t-1}} - \underset{(-3.252^{b})}{0.149p'_{t-5}} - \underset{(-1.963^{a})}{0.145p'_{t-11}} + \underset{(1.963^{a})}{0.100f'_{t-11}}$$

S.E.E. $= 0.007$ DW $= 1.977$ $F(4, 975) = 4.101^{b}$ $R^2 = 0.017$

Panel C: Second Sub-sample, 25/05/1989–28/04/1993

$$\begin{array}{llllllllllllllll}
p'_t & = & 0.000 & - & 0.00002Z^P_{t-1} & - & 0.186p'_{t-1} & + & 0.134p'_{t-5} & - & 0.064p'_{t-8} & + & 0.150f'_{t-1} & + & 0.089f'_{t-2} & - & 0.036f'_{t-6} \\
 & & (0.748) & & (-0.142) & & (-4.107^b) & & (-4.262^b) & & (-2.098^a) & & (6.118^b) & & (5.155^b) & & (-2.168^a) \\
\text{S.E.E.} & = & 0.0021 & & & & \text{DW} = 2.005 & & & & F(7,975) = 11.78^b & & & & R^2 = 0.078 & & \\
 \\
p'_t & = & 0.000 & + & 0.00000Z^F_{t-1} & - & 0.187p'_{t-1} & + & 0.134p'_{t-5} & - & 0.065f'_{t-8} & + & 0.150f'_{t-1} & + & 0.089f'_{t-2} & - & 0.036f'_{t-6} \\
 & & (0.747) & & (0.022) & & (-4.115^b) & & (4.262^b) & & (-2.096^a) & & (6.142^b) & & (5.167^b) & & (2.159^a) \\
\text{S.E.E.} & = & 0.0021 & & & & \text{DW} = 2.005 & & & & F(7,975) = 11.78^b & & & & R^2 = 0.078 & & \\
 \\
f'_t & = & 0.000 & - & 0.00125Z^F_{t-1} & + & 0.165p'_{t-1} & + & 0.180p'_{t-5} & & & & & & & & \\
 & & (0.361) & & (-4.969^b) & & (2.689^b) & & (3.014^b) & & & & & & & & \\
\text{S.E.E.} & = & 0.0040 & & & & \text{DW} = 1.996 & & & & F(3,982) = 12.09^b & & & & R^2 = 0.036 & & \\
 \\
f'_t & = & 0.000 & + & 0.00017Z^P_{t-1} & + & 0.157p'_{t-1} & + & 0.175p'_{t-5} & & & & & & & & \\
 & & (0.368) & & (4.653^b) & & (2.566^a) & & (2.922^b) & & & & & & & & \\
\text{S.E.E.} & = & 0.0040 & & & & \text{DW} = 1.998 & & & & F(3,982) = 11.06^b & & & & R^2 = 0.033 & &
\end{array}$$

Note: Z^P_{t-1} and Z^F_{t-1} are the residuals from regression eq. (3) when P_t and F_t are the LHS variable, correspondingly.

[a]Indicates the test statistic is significant at the 5% level.

[b]Indicates the test statistic is significant at the 1% level.

equilibrium error and lagged innovations in spot and futures series; (c) index futures innovations are regressed on last period's futures (own-market) equilibrium error and lagged innovations in spot and futures series; (d) index futures innovations are regressed on last period's index (cross-market) equilibrium error and lagged innovations in spot and futures series. If the cointegrated variable, X_t, adjusts towards the long-run equilibrium, a priori, the error correction coefficient is expected to be negative for the own-market equilibrium error term and positive for the cross-market equilibrium error term.

The model-building strategy advocated in Engle and Granger (1987) is adopted. First, a preliminary model which includes 12 lagged p_t' and f_t' is employed.[12] The final ECM only includes those lagged terms with statistically significant regression coefficients. The numbers in parentheses under the error correction equations are the associated t-statistics. The standard error of estimate (S.E.E.), R^2, Durbin–Watson, and F-statistics are also reported. The full sample results are presented in Panel A. It can be seen that the OLS regression coefficients for error correction terms Z_{t-1}^P and Z_{t-1}^F have the expected sign and are statistically significant, which suggest that both the index spot and futures prices adjust towards the long-run equilibrium relationship when disequilibrium occurs. The coefficients for some lagged innovations and lagged cross-market innovations are also significant. Therefore, current innovations in both the index and index futures prices are attributable also to some lagged changes in the index and index futures prices. The results suggest that there are two-way feedback relationships in the two markets. It is noted that the t-statistics for the f_{t-1}' term is much higher than for any other lagged variables in the spot ECM equation. This suggests that the information flow from the index futures market to the index cash market is stronger. The low regression R^2 values are not surprising because most of the price movements should be contemporaneous if the markets are relatively efficient and frictionless. The ECM explains 8.0% of the variations in the index spot innovations and only 1.1% of the variations in index futures innovations. The relatively low R^2 for the futures ECM equations suggests that the index futures market plays a leading role in incorporating new information because lagged variables of both markets can explain only a very small portion of the current innovations. The results are qualitatively similar when the cross-market equilibrium errors are used in the ECM equations.

[12] It is found that p_{t-i}' and f_{t-j}' are not significant when $i,\ j \geq 12$.

The whole sample period is divided into two equal sub-samples to examine whether the cointegration relation is stable for the whole period. Panel B of Table VI presents the OLS of the error correction models for the first sub-sample period. For the index spot innovation regression, the results are similar to the full sample period. However, for the index futures innovation regression, the error correction coefficients are not statistically significant although they have the expected sign. This indicates that most of the adjustment is made through the cash market when disequilibrium occurs. Interestingly, for the second sub-sample, the error correction coefficients in the index spot innovation regression becomes insignificant, which suggests that most of the adjustment is now made through the futures market when disequilibrium occurs. These seemingly conflicting results are not surprising because as the MBI futures market matures, it becomes more cost effective to trade in the futures markets rather than the cash market. Therefore, the information discovery role of the futures market strengthens over the sample period.

Model Diagnostics

The OLS regression method assumes that the disturbance term, v_t, follows an independent, identical, and normal distribution. The residuals from the ECM eq. (7) are examined for statistical adequacy. It is found that the serial correlations for the residuals are small and insignificant. However, there is strong evidence of heteroscedasticity. The autocorrelation coefficients for squared residuals from the index and the index futures ECM equations are significant at the 1% level. The Ljung–Box Q(6) statistics are 720.8 and 382.7, respectively. Although the OLS estimates are consistent when the error term is heteroskedastistic, the standard errors of the estimates are biased. The t-statistics reported in Table VI may be invalid. Therefore, a heteroscedasticity-consistent estimation procedure proposed in White (1980) is used to reestimate the ECM equations. The results are reported in Table VII.[13] As expected, the regression coefficients are the same. However, the t-statistics are over-estimated, in general, by the OLS method. Some of the originally lagged innovation terms become marginally significant at the 10% level. Nevertheless, the qualitative interpretations remain similar to those of the OLS estimation results. It is interesting to note that although none of the lagged innovations are significant at the 5% level in the futures

[13]Sub-sample results are similar to those of the full sample. They are available upon request.

TABLE VII

Heteroscedasticity-Consistent Estimation Results of Error Correction Models (Full Sample, 11/06/1985–28/04/1993)

p'_t	$= 0.000$	$- 0.00031Z^P_{t-1}$	$- 0.278p'_{t-1}$	$- 0.125p'_{t-5}$	$+ 0.226f'_{t-1}$	$+ 0.060f'_{t-2}$	$+ 0.058f'_{t-3}$	$+ 0.050f'_{t-5}$	$+ 0.051f'_{t-12}$
	(0.468)	(-2.042^a)	(-2.855^b)	(-2.405^a)	(5.063^b)	(2.298^a)	(2.133^a)	(1.739)	(2.415^a)
S.E.	$= 0.0039$		DW = 2.009		$F(8, 1961) = 21.37^b$		$R^2 = 0.080$		
p'_t	$= 0.000$	$+ 0.00025Z^F_{t-1}$	$- 0.279p'_{t-1}$	$- 0.126p'_{t-5}$	$+ 0.228f'_{t-1}$	$+ 0.061f'_{t-2}$	$+ 0.058f'_{t-3}$	$+ 0.051f'_{t-5}$	$+ 0.051f'_{t-12}$
	(0.467)	(-1.696)	(-2.858^b)	(-2.423^a)	(5.052^b)	(2.316^a)	(2.140^b)	(1.751)	(2.403^a)
S.E.	$= 0.0037$		DW = 2.010		$F(8, 1961) = 21.18^b$		$R^2 = 0.080$		
f'_t	$= 0.000$	$- 0.00052Z^F_{t-1}$	$- 0.098p'_{t-5}$	$- 0.124p'_{t-11}$	$+ 0.071f'_{t-11}$				
	(0.694)	(-2.423^a)	(-1.802)	(-1.783)	(1.559)				
S.E.	$= 0.0058$		DW = 1.961		$F(4, 1966) = 5.59^b$		$R^2 = 0.011$		
f'_t	$= 0.000$	$+ 0.00048Z^P_{t-1}$	$- 0.099p'_{t-5}$	$- 0.124p'_{t-11}$	$+ 0.070f'_{t-11}$				
	(0.688)	(-2.162^a)	(-1.822)	(-1.785)	(1.543)				
S.E.	$= 0.0059$		DW = 1.966		$F(4, 1966) = 5.08^b$		$R^2 = 0.010$		

Note: Z^P_{t-1} and Z^F_{t-1} are the residuals from regression eq. (3) when P_t and F_t are the LHS variable, correspondingly.
[a]Indicates the test statistic is significant at the 5% level.
[b]Indicates the test statistic is significant at the 1% level.

innovations regression, they are significant at the 10% level. This further indicates that the futures market plays a major role in discovering new information.

CONCLUSIONS

It is found that the MBI and the index futures prices are unit-root nonstationary series. Cointegration analysis reveals that there is a long-run equilibrium relationship between the MBI and the index futures. The error correction model is then used to investigate the short-run dynamics and the price movements in the two markets. The results suggest that there is a two-way feedback relationship between the cash and futures markets and that most of the price movements are contemporaneous. However, the information flow from the futures market to the cash market is stronger. Sub-sample results also suggest that for the first sub-sample period, the cash market plays a major role in adjusting towards a long-run equilibrium pricing relationship; whereas, for the second sub-sample period, the futures market plays a predominant role in eliminating the equilibrium errors.

These findings are useful to investors in the municipal bond market. Investors can potentially improve hedging performance by incorporating the long-run cointegrating relationship and short-term dynamics between the MBI and the index futures prices.

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