The Long-Run Linkage Between Yields on Treasury and Municipal Bonds and the 1986 Tax Act

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Abstract

The Tax Act of 1986 changed the tax treatment of tax-exempt municipal bonds for banks. Since banks were the dominant participant in the municipal bond market until 1986, some believe that this resulted in a breakdown of the long-run equilibrium relationship between municipal and U.S. Treasury securities of equal maturity. We present evidence that there was a significant structural break in the relationship between municipal and Treasury bonds around the time of the Tax Act but that once the break is accounted for, the relationship remains intact.

1 Introduction

The relationship between tax-exempt municipal bonds and taxable U.S. Treasury bonds of the same maturity should be of an equilibrium nature, at least in the long run. If the yield on municipals (munis hereafter) rose high enough, holders of Treasury securities would be induced to sell and buy munis with the proceeds. The simultaneous selling of Treasury securities and buying of munis would drive the yield on the former up and the latter down until the new equilibrium between them was reached. This suggests that the two series, Treasuries and munis, cannot drift too far apart without an equilibrating market reaction.[1]

Until 1985 the long-run relationship between Treasuies and munis seemed to obey the simple equilibrium described above. However, around 1985 the linkage between municipal and Treasury yields changed. In particular, the spread between the two yields decreased at all maturities.

The Tax Act of 1986 played an important role in decreasing the slope of the municipal yield curve as it dramatically altered the tax treatment of municipal investments for commercial banks. Prior to 1986 commercial banks were the major investor in the municipal bond market and thus were the representative investor. During the mid-1980s the deductibility of interest paid to carry tax-exempt bonds was eliminated. Banks lost their incentive to invest in municipal bonds, and the dominance of the municipal bond market shifted from commercial banks to individuals, primarily through investments in mutual funds. Households have now become the largest holder of tax-exempt debt in the years following 1986. Individual investors subject to lower marginal tax rates have now replaced banks, but require higher relative yields to attract them to the municipal bond market. Thus, municipal bond rates, particularly those of shorter maturity such as one-year, have increased relative to Treasury security yields.

The purpose of this paper is to investigate whether a long-run stable linkage exists between municipal bonds and Treasury securities. Specifically, we examine whether the tax changes of 1986 caused a permanent shift in the long-run relation between these two securities or simply a level shift in the trend (mean), leaving the long-run linkage intact. We also investigate the linkage between municipal bonds and inflation and find that a one-to-one correspondence exists between 1-year municipal bond yields and 1-year

[1]This assumes that any risk premia, while time varying, are stationary.

inflation rates. These results show that any structural change in the relationship between munis and Treasuries is not due to changes in the inflation regime. The unitary Fisher effect in the municipal bond market also provides evidence in favor of the Darby (1975) hypothesis.

A limitation of the early empirical literature investigating the linkages between Municipal bond yields and Treasury yields was that it did not utilize appropriate econometric procedures to account for the nonstationarity of municipal and Treasury bond yields. This aspect meant that this literature was not able to distinguish between the short-run and the long-run. This study attempts to delineate more clearly the short-run and long-run between Treasury bill rates and municipal bond rates. In this way this paper corrects the mis- specifications in the earlier literature by examining the stationarity of municipal bond rates, Treasury bill rates and inflation rates, modeling short-run and long-run dynamics through the use of cointegration and using error-correction methodologies, as well as an examination of the stability of the long-run linkages between these variables.

The rest of the paper is organized as follows. Section 1 gives a short theoretical description of the relationship between Treasury and municipal bond yields and inflation rates. Section 2 discusses the effect of the 1986 Tax Act on the municipal bond market. Section 3 provides a description of the econometric methodology. Section 4 presents the empirical results and section 5 concludes with a discussion of the results within the context of the existing literature.

2 The Linkage Between Municipal and Treasury Bonds

Many prior studies have examined interrelationships between bond yields. Frequently they have employed the Miller (1977) hypothesis that debt levels are irrelevant and bond yield differentials are due to such effects as different tax features of Treasury and municipal bonds. For example, Trzcinka (1982) finds that municipal bonds consistently have been approximately 52 percent of the corresponding corporate bond yield, as predicted given the prevailing tax rate in his sample. An alternative hypothesis that municipal yields are determined by market segmentation has received less than unanimous empirical support. For example, Kidwell and Koch (1982, 1983) and Kochin and Parks (1988) contend that market segmentation exists while Campbell (1980) argues that it does not.

This section describes the traditional model of municipal market equilibrium. Previous studies use a simple model relating the return on a municipal bond to that on a taxable bond. We consider a taxable investor with a marginal tax rate τ. If the investor is to be indifferent between holding a par taxable bond having a yield of i^T and a par tax-exempt municipal bond having a yield i^M, the following relation should hold:

$$i^T = \left[\frac{1}{1-\tau}\right] i^M \tag{1}$$

Bierman and Hass (1975), Kidwell and Trzcinka (1982) and Yawitz, Maloney, and Ederington (1985) extend this framework.[2]

A key question in the analysis of the linkage between Treasury and Municipal bonds is who is the representative investor? That is, whose tax rates matter? Fama (1977) developed a model in which he argued that corporations (particularly commercial banks) were the marginal investor in all short-term municipal securities, and hence the pricing of these securities depended on the tax rates which these banks faced. Because they can borrow at the corporate bond rate and deduct their interest payments, the after-tax cost of funds is $(1-\tau_c)i^T$ where τ_c is the corporate tax rate and i^T is the taxable Treasury bond rate. If the opportunity to invest in municipal bonds is unrestricted, commercial banks will hold municipal bonds when the yield on municipal bonds, i^M, is greater than $(1-\tau_c)i^T$ and hold taxable bonds when i^M is less than $(1-\tau_c)i^T$. Security market arbitrage will ensure that $i^M = (1-\tau_c)i^T$ in equilibrium as banks borrow at the taxable rate and invest at the tax-exempt rate.

The early empirical evidence supported Fama's (1977) view that; i) personal income taxes do not matter in the determination of municipal bond yields and ii) that the implicit tax rate is determined by the corporate income tax rate.[3] Trczinka (1982) found that the implicit tax rate was close

[2]For example, Yawitz, Maloney and Ederington (1985) assume default with no partial payment.

[3]The yield differential between taxable Treasury bills and tax-exempt municipal bonds is often summarized by the implicit tax rate that equates the after-tax yield from a taxable and a tax-exempt security. It is also called the break-even tax rate because an investor with that marginal tax rate will be indifferent between a tax-exempt bond and a taxable bond. This implicit tax rate is given by $\tau^* = \frac{i^T - i^M}{i^T}$.

to the corporate income tax rate during the 1970s. However, later evidence has led to the rejection of Fama's view. For example, Fortune (1988) showed that Trczinka's results were unique to the 1970s and that during the 1980s the movements in the implicit tax rate were consistent with changes in personal income tax rates. Such a shift is consistent with changes in the tax code during the 1980s which weakened Fama's view. Fortune (1996) shows that the implicit tax rate for 5- and 20-year maturity municipal bonds and Treasury bonds fell sharply in 1981, when the Economic Recovery and Tax Act (ERTA) reduced marginal tax rates of the wealthy but did not change the corporate tax rate. This event occurred before the 1986 Tax Act, which shifted the dominance of the municipal bond market from commercial banks to individuals who post-1986 appeared to be the marginal investor. Such evidence suggests further negative evidence against Fama's view even prior to 1986. Thus, it is still a debate as to who should be designated as the representative investor of municipal bonds and which marginal tax rate is appropriate.

It is interesting to note how the value of the coefficient $\left[\frac{1}{1-\tau}\right]$ in equation (1) linking Treasury bond yields and municipal bond yields would differ depending on whether one uses the personal income tax rate or the corporate income tax rate. The values of the highest corporate tax rate in effect for some selected periods are given as follows: 1965-1967: τ_c =0.48; 1968-1969: τ_c =0.528; 1970: τ_c =0.492; 1971-1978: τ_c =0.48; 1979-1986: τ_c =0.46; 1987-1992: τ_c =0.34. If we assume an average corporate tax rate of 0.49 prior to 1986, then the value of the coefficient in equation (1) would be 1.96. Given average marginal tax rates faced by households in the United States over the period 1950 to 1995, the coefficient linking Treasury bond yields to municipal bond yields should lie somewhere in the range of 1.30 to 1.50.[4]

Another important linkage is the Fisher relation. The Fisher equation encapsulates the simple relationship hypothesized to exist between nominal interest rates and expected inflation first delineated by Irving Fisher (1930)[**?**]. If the ex-ante real rate of interest is assumed constant, then self-interested economic agents will require a nominal return that not only compensates for the marginal utility of foregone current consumption (measured by the real interest rate), but a nominal return that compensates for the decline in the purchasing power of money over the term of the loan. The decline in the

[4]Barro and Sahasakul (1986) compute an average marginal tax rate of 0.30 over this period.

purchasing power of money is commonly proxied by the price inflation that is expected to occur over the life of the loan.

When nominal returns are subject to taxation, the tax-adjusted "observable" Fisher equation is given in equation (2),

$$i_t^T = \left[\frac{1}{1-\tau}\right]\pi_{t+1} + \left[\frac{1}{1-\tau}\right] r_t + \varepsilon_{t+1} \tag{2}$$

where τ is the average marginal tax rate. Equation (2) is derived by noting that when nominal interest is taxed at rate τ, the after-tax nominal return is $i_t^T\left[1-\tau\right]$ in equation (2). Equation (2) implies a Fisher effect greater than one for all tax rates greater than zero. The interest income derived from U.S. Treasury securities is subject to ordinary income tax and equation (2) represents the appropriate Fisher relation.[5] The municipal securities examined in this paper are free from federal income tax, implying an alternative form of the Fisher equation for this type of asset given in equation (3) below.

For non-taxable securities, the Fisher equation is given in its most simple form as,

$$i_t^M = \pi_{t+1} + r_t + \varepsilon_{t+1} \tag{3}$$

where i_t^M is the nominal interest rate on municipal bonds, r_t is the real interest rate, π_{t+1} is the inflation rate from period t to t+1 and $\varepsilon_{t+1}$is a rational expectations forecast error.[6] Equation (3) demonstrates that changes in inflation should be reflected by equal changes in tax-free nominal interest rates when the real rate is assumed to be constant. The response of nominal interest rates to (expected) inflation has been called the "Fisher effect". If economic agents do not suffer from what Tanzi (1980) calls "fiscal illusion" then we should observe that Fisher effect estimates from municipal securities are significantly smaller than analogous estimates from Treasury securities. Specifically, if Tobin (1965,1969) effects and changing inflation dynamics are short-run in nature, as implied by the long-run superneutrality hypothesis, the estimate of the municipal bond Fisher effect should be insignificantly different from one while that based upon Treasury securities should be greater than one.

[5]This assumes that the securities are not held in a qualified tax-deferred annuity or similar vehicle.

[6]When economic agents are uncertain about their future consumption path there will also be a risk premium term in the Fisher equation (4). Smith (1993) and Ireland (1996) demonstrate that in the U.S. this risk premium term is negligible.

3 The 1986 Tax Act and the Municipal Bond Market

Prior to 1986, Commercial banks were in a privileged position. They could borrow funds (say through issuing deposits) and then deduct the interest on these borrowings as long as they invested the funds in municipal securities. Congress recognized that banks had the capability of arbitraging the differential between their after-tax cost of borrowing and tax- exempt municipal yields, and in 1983 disallowed the deduction of 15 percent of bank municipal carrying costs associated with new municipal purchases. In 1985, Congress increased the lost deduction to 20 percent. Despite this, banks were still able to deduct the majority of their interest expenses. Effective after-tax municipal yields were still greater than banks' after-tax cost of borrowing, and municipal bonds remained attractive investments to commercial banks.

The Tax Act of 1986 cut back dramatically on commercial bank purchases of municipal securities. This Act designated two classes of securities: bank-qualified and non- qualified municipal securities.[7] Currently, commercial banks are allowed to deduct 80 percent of carrying costs associated with the purchase of bank-qualified municipal securities. The deduction for carrying costs for non-qualified municipal securities has been eliminated completely. When commercial banks lost the entire interest deduction, the additional tax- liability offset most of the benefits that banks had gained from receiving tax-exempt municipal interest. Thus, the effective after-tax return on a non-qualified municipal security is far below that of any comparable-maturity Treasury security. The effect of the Tax Act of 1986 then was to remove banks as investors in non-qualified municipal securities. On a risk- adjusted basis any comparable-maturity bank-qualified municipal or taxable security will provide a higher after tax-yield. Banks continue to find that bank-qualified municipal securities offer the highest after-tax yields. However, due to the fact that so few municipal bonds meet the small-issue, public purpose exception, and the costs of searching out these issues are high, banks have invested very little in bank-qualified municipal securities.

The ownership structure of tax-exempt debt is very different today than it was pre- 1986. Prior to 1986 commercial banks were the major investor in

[7] For a security to be classified as bank-qualified, as state or local government can issue no more than $10 million in securities per year and the proceeds must be used for an "essential " public purpose.

the municipal bond market and thus were the representative investor. During the mid-1980s the deductibility of interest paid to carry tax-exempt bonds was eliminated. Banks lost their incentive to invest in municipal bonds, and the dominance of the municipal bond market shifted from commercial banks to individuals, primarily through investments in mutual funds. Households have now become the largest holder of tax-exempt debt in the years following 1986. Farinella and Koch (1994) provides data on net purchases of municipal bonds by various investor groups over different time periods between 1960 and 1992. Commercial banks purchased the majority of net new municipal issues during the decades of the 1960s and 1970s (62 percent and 36 percent of the total respectively). From the period 1980-1986 individual investors replaced commercial banks at the top by purchasing 71 percent of net new issues. Banks continued to add to their portfolios but at a slower rate. This was because profits were low and the relative yield advantage of municipal bonds had decreased with the loss of interest deductibility. According to Farinella and Koch (1994), since the middle of 1986 commercial banks have been net sellers of municipal securities in each year, to a point where they held only 8 percent of total outstanding issues at the end of 1992. In contrast to this, in 1992 individuals held 75 percent either directly or indirectly through mutual funds.

4 Econometric Methodology

The estimation procedure employed in this paper is based upon the vector error correction model (VECM) in equation (4),

$$\Delta X_t = \mu + \alpha\beta' X_{t-1} + \sum_{j=1}^{k-1} \Gamma_j \Delta X_{t-j} + \xi_t \tag{4}$$

where X_t is a $p \times 1$ vector of possibly I(1) series, β is a $p \times r$ matrix whose r columns represent the cointegrating vectors among the variables in X_t and α is a $p \times r$ matrix whose p rows represent the error correction coefficients. The Johansen (1988, 1991)[**?**][**?**] test for cointegration, called the trace test, tests the rank of the $p \times p$ product matrix such that reduced rank, rank less than p, implies cointegration. The term μ captures the deterministic components in the series that are eliminated by the cointegration vector(s). Alternative assumptions about the nature of the deterministic components are discussed in detail in Johansen (1994). Johansen describes five different cases. These distinctions are important since the asymptotic distribution of the test statistics for cointegration depend upon the specification of the deterministic components as demonstrated in Johansen (1994).

In the most general case, the trend specification allows for quadratic trends in the levels of the data and linear trends in the error correction representation. In this very general case these linear trends exist in both the differenced VAR model as well as in the cointegrating equations. Both the differenced VAR and the cointegrating equations are characterized by nonzero intercepts. In the next, restricted case, the same general specification of the intercepts remains but the linear trends are confined to the cointegrating equations, that is the cointegration vectors are trend stationary. The next restricted case, which we will call case number 3, is the most common speciofication in the literature. In case 3 there may be linear deterministic trends in the levels of the data but these are eliminated by the cointegrating relation. In the next restrictive case, case 2, the intercept is confined to be in the cointegrating equation implying that there are no deterministic trends in data. The most restrictive case, case 1, restricts all deterministic components to be zero, except for possible intervention dummy variables. Johansen (1994) demonstrates how these alternative trend specifications have important influences on the likelihood based test for cointegrating rank. Furthermore, the inclusion of a superfolous non-zero mean into the specification can bias estimates of the cointegrating coefficients. Since this paper investigates interest rates and inflation, which have no trends, we limit our attention to cases 3, 2, and 1.

5 Empirical Analysis

The standard source of information on the yield curve for both taxable and tax-exempt securities is Salomon Brothers' Analytical Record of Yields and Yield Spreads. This source provides estimates of the yield curve for par bonds on the first of each month. This study employs monthly data, over the period January 1950 to December 1995, on i) 1-year Treasury bill yields, ii) Consumer Price Index, iii) 1-year prime grade municipal bond yields. We focus on 'Prime' grade municipal bonds as these are the least risky category of municipal bonds and are thus most comparable to Treasury securities. The Treasury bill data and the municipal bond data are from Salomon Brothers.

The consumer price index is from Citibase.[8] Annualized percentage changes in the CPI are used as proxies for expected inflation. We focus our attention on short-term securities with 1-year maturity for a number of reasons. First, commercial banks primarily invested in short-term municipal bonds during the pre-1986 period. Second, longer term yields are influenced by factors difficult to control for, and thus, make it difficult to compare Treasury bonds with municipal bonds with long maturities. Fortune (1996) has pointed out a number of factors which contaminate long-term yields and make US Treasury bonds and municipal bonds less comparable at longer maturities.[9] Furthermore, since we were also interested in whether the Tax Act of 1986 altered the relation between municipal bond yields and inflation we needed to use short-term maturities to match up with estimates of inflation.

The top panel of figure 1 plots the one-year Treasury bill interest rate, the middle panel shows the one-year prime grade municipal bond yield and the bottom panel displays the annualized CPI inflation rate. Standard augmented Dickey-Fuller (ADF) t-tests are presented in table 1. Using the critical value for this test given by Dickey and Fuller (1976)[**?**] of -2.79, the null hypothesis of a unit root cannot be rejected for any of the three series when a lag of 6 or larger is used.[10]

The non-stationarity of the data and the assumption that Tobin effects are negligible in the long run implies that the use of cointegration techniques is most appropriate in analyzing the two linkages represented in equation (1) and (3). The discussion in section I suggests that the cointegrating vectors should be of the form $[1, -\beta_{TM}]'$, where $\beta_{TM} > 1$ for the linkage between Treasury bill and municipal bond yields (eq. 1) and $[1, -\beta_{\pi}]'$, where $\beta_{\pi} = 1$, for the relation between municipal bond rates and inflation rates. If one assumes an average corporate income (individual marginal) tax rate of 0.49

[8] The CPI series we use is PUNEW. The results are not sensitive to alternative choices for a price series.

[9] These factors include; i) municipal bonds and U.S. Treasury bonds with the same term to maturitynecessarily differ in their duration and, therefore, in the interest rate risk their holders experience, ii) the fact that U.S. Treasury bonds do not carry call features while many municipal bonds do, and iii) the default risk of U.S. Treasury bonds is virtually zero, while that of municipal bonds with maturities of five years or more may be non-negligble.

[10] It is well documented that post-war U.S. inflation has a large negative moving average component that creates large size distortions in the ADF tests, see Crowder and Hoffman (1996)[**?**], Ball and Cecchetti (1990), and Crowder (1996). Said and Dickey (1984)[**?**] and Schwert (1987)[**?**] suggest using high order autoregressive polynomials to approximate the MA component.

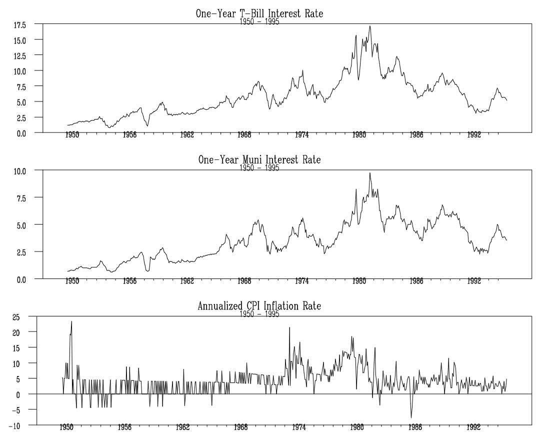


Figure 1: Data

(0.30) the cointegrating vector between the 1-year Treasury bill rate (inflation rate) and the 1-year municipal bond rate should be $[1, -1.96]'$ ($[1, -1.42]'$).

The discussion above suggests that the appropriate tax rate in this context is open to debate, but lies somewhere between 0.30 and 0.49, thus the cointegration coefficient linking Treasury bonds and municipal bonds should lie somewhere between 1.42 and 1.96. In addition, the cointegrating vector between municipal bond rates and inflation should be $[1, -1]'$.

These hypotheses are tested using the full information maximum likelihood (FIML) procedure proposed by Johansen (1988, 1991). The results of the cointegration analysis between Treasury bill yields and municipal bond yields are reported in table 2. The treatment of the deterministic components follows the discussion above.[11]

[11] Case 3 $H_1(r)$ denotes the specification that allows for a deterministic trend in the data that is eliminated by the cointegrating relation. The critical values appropriate for this case can be found in table 1 of Osterwald-Lenum (1992)[**?**]. Case 2 $H_{1*}(r)$ restricts the data to have no trends but allows for a non-zero constant in the cointegrating vector. Critical values for this case are taken from Osterwald-Lenum (1992) table 1*. The final and most restrictive case 1 $H_0(r)$ specifies the cointegrating vector to have a zero mean. Critical values appropriate for this specification are from Osterwald-Lenum (1992) table

Panel A of table 2 presents the tests for cointegration between the one-year municipal bond yield and the one-year U.S. Treasury bill yield. The results suggest that the two series are not cointegrated or that they do not share a long-run equilibrium relationship. This is true regardless of the specification of the deterministic components. Note also that two of the three estimates of the cointegrating parameter are outside of the range implied by the average marginal tax rates in the U.S. of the sample period.

Conventional tests for cointegration are usually conducted under the assumption that the long-run equilibrium between the variables in question is stable over the entire sample period. This may be a misleading assumption, especially if there were significant changes in the economic environment over the sample period. As the theoretical foundation of this paper suggests, there was a potential change in the linkage between the yields on Treasury bonds and municipal bonds about 1986. Previous studies have also documented a change in the inflation process during the money targeting period of the Fed from 1979:11 - 1982:10. We conduct dynamic econometric specification analysis to investigate the existence of such regime changes and account for them in the estimation results.

In this regard, Hansen and Johansen (1993) have developed a recursive likelihood ratio (LR) test for the constancy of the cointegrating space within the FIML estimation procedure. This procedure fixes the short-run dynamics at the full sample estimates and treats the full sample estimate of the cointegration vector as the null hypothesis in the recursive tests. This implies that rejections of stability are due to a change in the long-run relationship not due to shifting dynamics. Recursive LR tests for the stability of the cointegrating vector between munis and Treasuries over the entire sample period are displayed in figure 2. The recursive estimates of LR statistics have been normalized by the appropriate 5% critical value so that values greater than one imply statistical significance. The results suggest instability in the cointegration vector occurring in the early 1980s. We hypothesize that this structural break is in anticipation of the effects of the Tax Act of 1986. Since it is believed that this break represents an exogenous change in the municipal-Treasury market, a dummy variables is used to capture this effect in the VECM. The dummy variable takes on a value of zero until 1985:1 when it takes on a value of one until the end of the sample.[12] Now all of the

0. We investigate trend reduction through the use of likelihood ratio (LR) tests that allow one to test for the appropriate deterministic component specification for the VECM.

[12]Although the Tax Act of 1986 became effective in August 1986, municipal bond rates

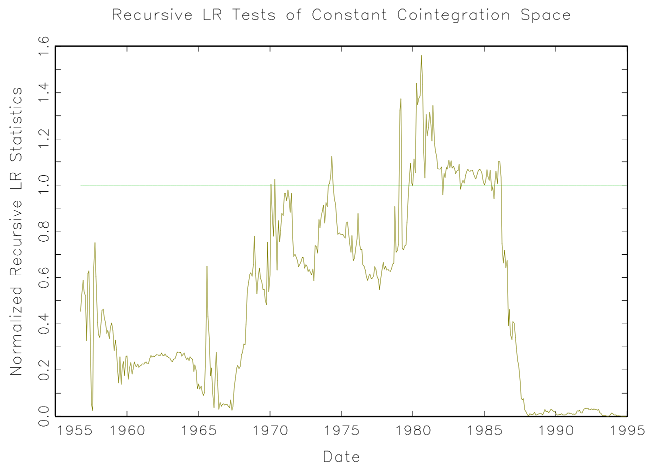


Figure 2: Recursive Test for Constant Cointegration Vector between Municiapl Yield and Treasury Yield- No Regime Shifts

specifications imply significant cointegration between munis and T-bills. The estimate of the cointegrating relation with the regime shift dummy included is within the theoretically implied interval. The estimate that includes a constant in the cointegrating vector is equal to -1.90 and has a standard error of 0.29.

Figure 3 presents the recursive LR tests of the constant cointegration space when the regime shift is included. There is now no evidence of any statistically significant break in the long-run equilibrium between municipal and T-bill yields. The one shift in 1985 captured all of the instability in the long-run relationship.

The marginal investors in municipal bonds prior to 1985 were commercial banks who invested primarily in short-term municipal bonds. Post-1985 the marginal investors were households who tended to invest in longer term

began to rise in the middle of 1985, as banks began withdrawing in large numbers from the municipal bond market. No exact month can be identified as the date in which a regime change actually began. The timing of an economic 'event' is notoriously difficult to identify because the real question is not when did legislation pass, or when was it proposed, but when did expectations about future consequences of the event take affect.

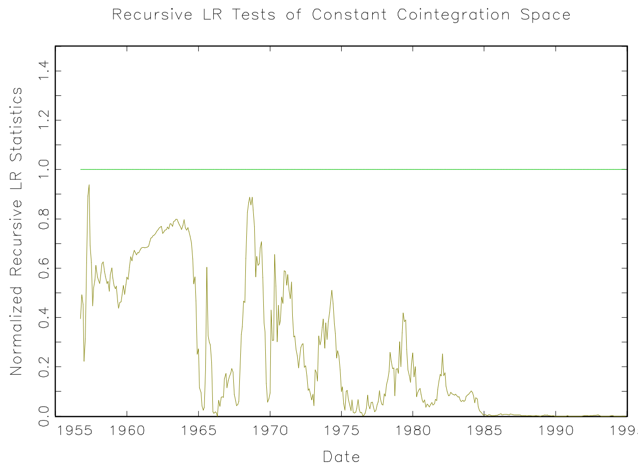


Figure 3: Recursive Test of Constant Cointegration Vector between Municipal Yield and Treasury Yield- Corrected for Regime Shift in 1985

municipal bonds. This suggests that the equilibrium parameter that determines the long-run relationship between munis and Treasuries may be different in the two periods. Since corporate average marginal tax rates were approximately 49% in the period prior to the Tax Act, the period appropriate for treating banks as the marginal investor, we may expect to find a cointegration vector estimate over this period statistically not different from $[1, -1.96]'$. The cointegrating vector estimate between Treasury bill rates and municipal bond rates for the period 1950:1-1985:12 is $[1, -1.79]'$ with a standard error on the normalized cointegration parameter of 0.21. The estimate is within two standard errors of the hypothesized value. Similarly, when we examine the post-1985 period, one in which households, with an average marginal tax rate of 30%, became the marginal investor, we obtain an estimate of $[1, -1.39]'$. The estimated standard error on the normalized cointegrating parameter is 0.15. This is also within two standard errors of the theoretical value.

Since households are the marginal investor in the post-1985 and hold primarily longer term municipal bonds, we investigate the relation between long-term Treasury bond yields and municipal bond yields for the post-1985 period. The cointegrating vector between Treasury bond and municipal bond yields for the period 1985:1-1995:12 for 5-year maturity bonds is $[1, -1.35]'$, $[1, -1.30]'$ for 10-year maturity bonds, and $[1, -1.19]'$ for 20-year maturity bonds. For the period 1950:1-1984:12, the period in which banks dominated the municipal bond market, the cointegrating vector estimates are $[1, -1.60]'$ for 5-year maturity bonds, $[1, -1.43]'$ for 10-year maturity bonds, and $[1, -1.28]'$ for 20-year maturity bonds. Regardless of the maturity, the cointegrating coefficient is smaller for the post-1985 period relative to the pre- 1985 period. When we test for cointegration over the full sample period, allowing for a level shift in 1985:1, we find the cointegrating vector between Treasury bond and municipal bond yields to be $[1, -1.61]'$ for 5-year maturities, $[1, -1.44]'$ for 10-year maturities, and $[1, -1.29]'$ for 20-year maturities. There is a clear inverse relationship between the value of the cointegrating coefficient and the term of the instruments examined. In addition, the earlier period seems to dominate the results.

The average implicit tax rate over the period 1950:1 to 1984:12 is 0.43 for 1-year maturities with a standard deviation of 0.07. The average implicit tax rate over the period 1985:1 to 1995:12 is 0.29 for 1-year maturities with a standard deviation of 0.06. This clearly shows how the spread between taxable and tax-exempt securities has changed after 1985. These values are comparable with the average highest corporate tax rate over the period 1965:1 - 1984:12 which was 0.49, and 0.34 for the period 1985:1-1995:12.

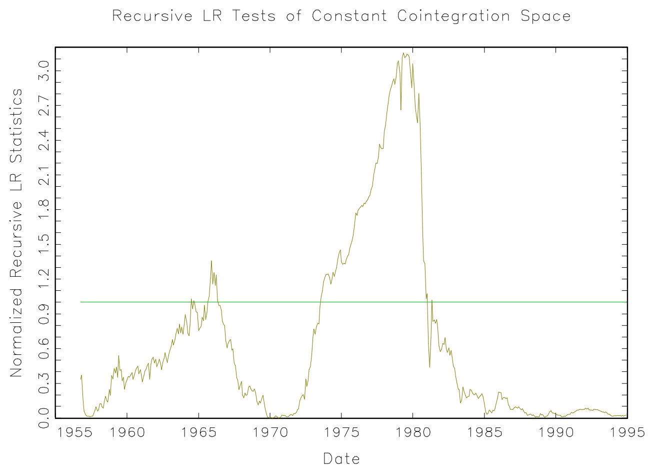
Why might the cointegrating vector estimate, which is $[1, -1.79]'$, during the pre-1986 period lie below the theoretical value of $[1, -1.96]'$ implied by an average corporate tax rate of 0.49. As discussed previously, prior to 1986 commercial banks could deduct all of their financing costs when investing in municipal bonds. With the Tax Reform Act of 1986, banks can only deduct 80 percent of their borrowing costs associated with buying municipal bonds issued for essential public purposes, if the municipality issues less than \$10 million in securities per year. Banks lost the entire interest deduction on financing costs for new purchases of all other municipal securities. This suggests that commercial banks received a form of subsidy or tax credit when they purchased municipal securities prior to 1986. This implies that equation (1) could be rewritten as equation (5) below:

$$i^T = \left[\frac{1}{1 - \tau + \omega}\right] i^M \tag{5}$$

where ω is a premium (or subsidy) which lowers the tax rate due to the interest deductibility attached to purchases of municipal securities. A positive premium would decrease the cointegrating coefficient from its upper theoretical value of 1.96. A cointegrating vector of $[1, -1.79]'$ suggests that this premium term, , takes on the value of 0.05, reducing to a value of 0.44, which is insignificant from the tax rate implied by the cointegration estimates of 0.42 during this period. Again, the evidence fits quite nicely with the hypothesized values.

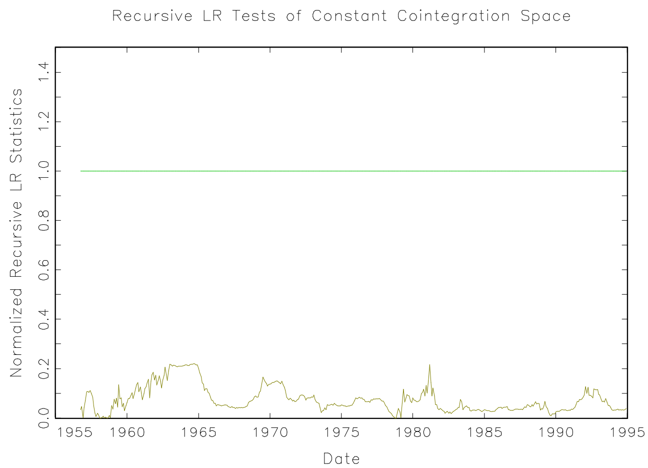
5.1 The Fisher Relation

Table 3 reports cointegration results for the linkage between the 1-year prime municipal bond rate and the 1-year annualized inflation rate. Panel A of table 3 reports results assuming no structural change over the sample period. These results indicate that we cannot reject the null of one cointegrating vector. Deterministic trend reduction leads to the conclusion of a zero mean in the cointegrating vector (case 1). The cointegrating vector between the 1-year prime municipal bond rate and the 1-year annualized inflation rate is $[1, -0.96]'$. This is insignificantly different from $[1, -1]'$. Stability tests using the LR recursive statistics are plotted in figure 4 and suggest two periods in which the municipal bond-inflation Fisher effect may not be stable. The first break or regime change occurred in the mid 1960s. The second break date appears to have occurred in the late 1970s through the early 1980s. This break date seems to correspond with the Federal Reserve operating policy change that occurred in October 1979. Since it is believed that this break represent an exogenous changes in the U.S. economy, dummy variables are used to capture their effects in the VECM. The first dummy variable takes on the value of zero until 1964:1 when it takes on a value of one through the end of the sample. The second dummy variable takes on a value of zero through 1965:12 and a value of one for the remainder of the sample. The third dummy variable take on a value of zero until 1979:11 and then one until the end of the sample. The fourth dummy variable takes on a value of zero through 1982:10 and one until the end of the sample. This allows for regime shifts during the periods 1964:1-1965:12 and 1979:11-1982:10.



Recursive Test of Constant Cointegration Vector between Municipal Yield and Inflation- No Regime Shifts

The estimate of the Fisher effect with the regime shift dummies included in the estimation is now $[1, -1.13]'$ for case 1 which is again insignificantly different from one. Regardless of whether we account for these regime shifts, the unitary Fisher effect between Treasury and municipal yields is established. The stability of this relationship is tested recursively and these results are displayed in figure 5. When the two regime breaks are modeled explicitly, the relationship between municipal bond yields and inflation appears to be stable. One final point to note is that the structural breaks in the muni-Treasury relationship are unrelated to those in the muni- inflation relationship. This gives us some confidence that the regime shift between munis and Treasuries identified above is not due to a change in the inflation process that underlies both series.



Recursive Test of Constant Cointegration Vector between Municipal Yield and Inflation - Corrected for Regime Shifts in 1964-65 and 1979-82

6 Conclusion

This study has analyzed the long-run relationship between taxable U.S. Treasury yields and tax-exempt municipal yields over the period 1950:1 to 1995:12. The results suggest that a structural break in the relationship accurred around 1985. We believe this is due to the impending Tax Act of 1986 that altered the tax treatment for banks' purchases of municipal bonds. Once this break is accounted for in the estimation, the long-run relationship between the two yields is shown to be stable over the entire period.

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Table 1: ADF Tests

Lag Length	1-Year T-Bill	1-Year Municipal Bond	1-Year CPI Inflation
4	-1.84	-2.07	-3.26
6	-1.74	-1.97	-2.46
8	-1.67	-1.74	-2.00

NOTE: TEST STATISTICS ARE CALCULATED FROM THE REGRESSION $\Delta X_t = \alpha_0 + \alpha_1 X_{t-1} + \sum_{j=1}^{\ell} \gamma_j \Delta X_{t-j} + \varepsilon_t$, WHERE J IS THE LAG LENGTH REFERRED TO IN COLUMN ONE. 95% CRITICAL VALUE IS -2.79.

Table 2: Cointegration Analysis of One-Year Municipal and Treasury Yields

Panel A: No Regime Shifts			
Trace Statistics	Case3: $H_1(r)$	Case 2: $H_{1*}(r)$	Case 1: $H_0(r)$
r= 0	10.49	10.68	4.54
r$\leq$ 1	3.02	3.06	0.78
$\hat{\beta}$	-2.52 (0.81)	-2.50 (0.78)	-1.67 (0.41)
Panel B: Corrected for Regime Shift in 1985			
Trace Statistics	Case3: $H_1(r)$	Case 2: $H_{1*}(r)$	Case 1: $H_0(r)$
r= 0	20.43*	21.27*	14.58*
r$\leq$ 1	4.58	5.40	0.12
$\hat{\beta}$	-1.90 (0.27)	-1.90 (0.29)	-1.79 (0.21)

Notes: Case 3 allows for deterministic trends in the data. Critical values appropriate for case 3 can be found in table 1 of Osterwald-Lenum (1992). Case 2 restricts the data to have no deterministic trends but allows for the cointegration vector to have a non-zero mean. Critical values appropriate for case 2 can be found in table 1* of Osterwald-Lenum (1992). Case 1 restricts the mean of the cointegration vector to be zero. Critical values appropriate for case 1 can be found in table 0 of Osterwald-Lenum (1992). Lag length (k) in VAR = 13. No. of Observations = 552. Cointegration analysis in panel b account for a structural shift in mean of cointegrating vector after 1985:12. An * Denotes a significant statistic at the 95% level or higher. Numbers in parentheses are asymptotically valid standard errors.

Table 3: Cointegration Analysis of One-Year Municipal Yield and Inflation Rate

Panel A: No Regime Shifts			
Trace Statistics	Case3: $H_1(r)$	Case 2: $H_{1*}(r)$	Case 1: $H_0(r)$
r= 0	30.32*	31.45*	27.43*
r≤ 1	4.44	4.45	0.87
$\hat{\beta}$	-0.82 (0.24)	-0.82 (0.25)	-0.96 (0.13)
Panel B: Corrected for Regime Shifts in 1964-65 and 1979-82			
Trace Statistics	Case3: $H_1(r)$	Case 2: $H_{1*}(r)$	Case 1: $H_0(r)$
r= 0	50.21*	54.24*	44.82*
r≤ 1	12.88	13.20	7.47
$\hat{\beta}$	-0.59 (0.30)	-0.61 (0.30)	-1.13 (0.33)

Notes: Case 3 allows for deterministic trends in the data. Critical values appropriate for case 3 can be found in table 1 of Osterwald-Lenum (1992). Case 2 restricts the data to have no deterministic trends but allows for the cointegration vector to have a non-zero mean. Critical values appropriate for case 2 can be found in table 1* of Osterwald-Lenum (1992). Case 1 restricts the mean of the cointegration vector to be zero. Critical values appropriate for case 1 can be found in table 0 of Osterwald-Lenum (1992). Lag length (k) in VAR = 13. No. of Observations = 552. Cointegration analysis in panel b accounts for two regime shift dummy variables: 1964:1-1965:12 and 1979:11-1982:10. An * Denotes a significant statistic at the 95% level or higher. Numbers in parentheses are asymptotically valid standard errors.