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Are Tax Effects Important in the Long-Run Fisher Relationship? Evidence from the Municipal Bond Market

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ABSTRACT

Are nominal bonds appropriately discounted for taxes? Empirical estimates of the response of nominal interest rates to changes in inflation, the Fisher effect, have failed to produce a definitive answer. Four reasons have been put forward as possible explanations: (i) Tobin effects, (ii) fiscal illusion, (iii) peso problems, and (iv) different estimators. Utilizing data on taxable and tax-exempt bond interest rates and several different estimators, we find that the Fisher effect estimates are always larger for the taxable bond relative to the tax-exempt bond, suggesting that fiscal illusion and different estimators cannot account for the previous results.

HOW ECONOMIC AGENTS RESPOND to the tax burdens of assets with different tax treatment is an important issue at both the level of the individual (micro) and the aggregate economy (macro). One clear source of evidence would appear to be the market for high quality debt. Trading in both taxable and tax-exempt debt occurs on a daily basis. The response of the rates of return on taxable and tax-exempt debt to changes in (expected) inflation are described by the Fisher relationship.

The Fisher relationship is one of the oldest equilibrium relationships in financial economics. Fisher's original proposition was that if the real rate of interest is relatively constant over time, a fully perceived change in the purchasing power of money, as proxied by the inverse of the rate of inflation, would be accompanied by a proportional increase in the nominal interest rate. Darby (1975) points out that when nominal interest income is taxed, as are interest payments on U.S. Treasury bills, the Fisher relationship implies a response from nominal interest rates that is greater than the change in (expected) inflation in order to maintain the constant ex ante real interest rate. Summers (1983) suggests that given average marginal tax rates in the United States the appropriate value of the Fisher effect is 1.3 to 1.5. For tax-exempt bonds the original Fisher specification should be a good approximation implying a Fisher effect close to unity. In a seminal paper, Fama (1975) presents evidence to support the hypothesis that changes in nominal

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interest rates reflect changes in expected inflation. However, subsequent research contradicts most of Fama's original conclusions (see Nelson and Schwert (1977), inter alia). The evidence regarding the Fisher effect is inconclusive. Moreover, little light has been shed on the importance of taxes in determining the response of nominal interest rates to innovations in (expected) inflation. Though most of the recent studies cited utilize similar data over similar time periods, they each use a different estimator to generate point estimates of the Fisher effect. Most of the estimates of the response of nominal interest rates to changes in inflation, the Fisher effect, are much less than the value implied by theory.

There are four explanations offered in the literature for why empirical studies find low Fisher effect estimates. Taxable bonds may not fully respond to changes in (expected) inflation if economic agents suffer from some form of "fiscal illusion" (Tanzi (1980)) in which investors do not fully take into account the tax burden associated with a particular asset purchase. This would imply irrational behavior on the part of bond purchasers. It is entirely possible, however, to reconcile low Fisher effect estimates with economic agents behaving in a perfectly rational way. Tobin (1965, 1969) argues that agents shift out of nominal assets into real assets in response to increases in (expected) inflation. This causes the price of nominal assets to fall, thus increasing the expected return on them, and it causes the price of real assets to rise, thus reducing their expected return. The "Tobin effect" results in a negative relationship between inflation and the real rate of interest, thus depressing the Fisher effect.[1]

A third possible explanation for low Fisher effect estimates is offered by Evans and Lewis (1995), who hypothesize that regime switches in the U.S. inflation process may lead to estimates of the Fisher effect that are less than the implied theoretical value. This is the result of a "peso problem," in which a low probability is attached to a major event (e.g., high inflation), causing biased estimates. Ultimately the peso problem is a small sample problem. When sufficient data become available the peso problem should disappear if investors are truly rational. Finally, a fourth possible explanation for the low Fisher effect estimates found in empirical studies may be the result of using inappropriate estimators or misspecified estimation equations.[2]

These different explanations for the low Fisher effects cannot be disentangled by focusing only on taxable securities, as previous studies have done. In this paper we try to eliminate some of the possible candidates for low Fisher effect estimates by controlling for certain factors. First, we sidestep the issue of which estimator is most appropriate for the Fisher equation and

[1] Lucas (1980) argues that the Tobin effect should be a short-run phenomenon if monetary policy superneutrality holds. Thus, evidence of Tobin effects is indirect evidence against monetary superneutrality.

[2] Mishkin (1992) and Crowder and Hoffman (1996) are examples that stress using an appropriate estimator; Ng and Perron (1997) focus on proper specification of the regression equation within the context of estimating the Fisher effect.

focus instead on direct comparisons between Fisher effect estimates for Treasury bills and tax-free municipal bonds. If taxes are an important element in the determination of the relationship between nominal interest rates and inflation, then regardless of the estimator employed, one should consistently find that Fisher effect estimates on taxable bonds are greater than those on tax-free bonds.

We employ six estimators of the Fisher effect. These are ordinary least squares (OLS), Johansen's (1988) maximum likelihood estimator (MLE), Stock and Watson's (1993) dynamic OLS (DOLS), Engle and Yoo's (1991) three-step estimator (EY), Pesaran and Shin's (1995) autoregressive distributed lag (ARDL), and the fully-modified OLS (FM-OLS) estimator of Phillips and Hansen (1990). The results clearly demonstrate that, regardless of the estimator employed, taxes have a substantial influence on the size of the estimated Fisher effect. In particular, the Fisher effect estimate is always larger for taxable securities relative to tax-exempt securities. This finding provides evidence against fiscal illusion as a possible cause of low Fisher effect estimates.

The rest of the paper is organized as follows. Section I gives a short theoretical description of the relationship between nominal interest rates and inflation, Section II discusses the different estimators and presents the empirical results, and Section III concludes with a discussion of the results.

I. The Fisher Equation

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 mean of 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 purchasing power of money is commonly proxied by the price inflation that is expected to occur over the life of the loan. Standard asset pricing models in which both nominal and real bonds are traded yield the following condition:

$$i_t^k = r_t^k + E_t\pi_{t+k} + \tfrac{1}{2}\mathrm{Var}_t(\pi_{t+k}) - \mathrm{Cov}_t(\pi_{t+k}, q_{t+k}) \tag{1}$$

where i_t^k is the continuously compounded return on a nominal default-free k-period bond, r_t^k is the continuously compounded rate of return on a default-free bond that pays in real goods, π_{t+k} is the inflation rate from period t to $t+k$, q_{t+k} is the real stochastic discount factor, which is proportional to consumption growth in consumption based versions of the CAPM, and E_t, Var_t, and Cov_t are the expectation, variance, and covariance conditioned on information available in period t. Equation (1) relates the nominal interest rate to the sum of the real rate, expected inflation, and an inflation risk premium that depends on the second moments of π_{t+k} and q_{t+k} . If one assumes unbiased expectations and that the risk premium term is negligible, equation (1) can be written as in equation (2),

$$i_t^k = r_t^k + \pi_{t+k} + \varepsilon_{t+k} \tag{2}$$

where ε_{t+1} is a rational expectations forecast error and k is set equal to one.[3] In the long-run equilibrium or steady state r_t is constant and equal to r.[4] We interpret equation (2) to read that in the long run, changes in inflation should be reflected by equal changes in nominal interest rates. The response of nominal interest rates to (expected) inflation has been called *the Fisher effect*. Equation (2) implies a Fisher effect of one. However, when nominal returns are subject to taxation, the tax-adjusted "observable" Fisher equation is given in equation (3),

$$i_t^k = \left[\frac{1}{1-\tau}\right] r_t^k + \left[\frac{1}{1-\tau}\right] \pi_{t+k} + \left[\frac{1}{1-\tau}\right] \varepsilon_{t+k} \tag{3}$$

where τ is the average marginal tax rate. Equation (3) is derived by noting that when nominal interest is taxed at rate τ, the after-tax nominal return is $i_t(1-\tau)$ in equation (2). Equation (3) 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[5] so that equation (3) represents the appropriate form of the Fisher relationship. The interest income from municipal securities is free from most forms of taxation, implying that equation (2) is the relevant Fisher equation. 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.

II. Empirical Analysis

We employ several estimators to demonstrate the robustness of the relationship between the Fisher effect estimates from a taxable bond and the Fisher effect estimates from a tax-exempt bond. Three recent studies of the Fisher relationship (Mishkin (1992), Evans and Lewis (1995), and Crowder

[3] Ireland (1996) demonstrates that in the United States the inflation risk premium term is small enough to ignore. For our purposes it is only important that the risk premium have a constant mean over time.

[4] Rose (1988) demonstrates that the same asset pricing model that yields equation (1) implies that the expected real rate of interest is proportional to the real stochastic discount factor, which is itself constant in the steady state.

[5] This is true unless the securities are held in a qualified tax-deferred annuity or similar investment.

and Hoffman (1996)), produce very different estimates of the Fisher effect using very similar data. However, the studies use three different estimators, OLS, DOLS, and MLE, respectively. Since our purpose is not to determine which is the most appropriate estimator, but to determine whether taxes are important in explaining the response of interest rates to inflation, we employ six different estimators, the three used in prior analyses and three others, to estimate the Fisher effect associated with taxable and tax-exempt bonds.[6]

Each of the estimators is used to estimate β in the regression equation (4). Since the data appear to be nonstationary and the Fisher relationship is treated as a long-run equilibrium, one may interpret equation (4) as a cointegrating regression (see Engle and Granger (1987)),

$$i_t = \alpha + \beta\pi_{t+1} + e_t. \tag{4}$$

Clearly nominal interest rates and inflation are variables that are simultaneously determined within a larger economic system. Also, because the variables have unit roots or are nonstationary, the asymptotic distribution of the estimates is no longer Gaussian. Furthermore, the residual in equation (4), e_t, may be serially correlated. Though OLS estimates of β in equation (4) will be consistent, they may still suffer from biases induced by the simultaneity, unit roots, and serial correlation. Each of the other five estimators employed corrects for these potential sources of bias. It is uncertain, however, how well any of the five will perform in finite samples.

OLS, DOLS, FM, and ARDL are all single-equation estimators. Ng and Perron (1997) demonstrate that when the regressor is nearly stationary, as is the case with inflation, these estimators may deliver estimates that are severely biased. They suggest reversing the regressor/regressand in equation (4) in order to obtain better estimates. This amounts to putting the more integrated variable(s) on the right-hand side of the regression equation. The intuition for this suggestion is straightforward. Since the more integrated variable(s) has (have) greater variability, the parameter estimates are more precise.[7] MLE and EY are systems estimators so that the distinction between dependent and independent variables is meaningless.

Monthly data, over the period January 1950 to December 1995, are employed for (i) 1-year Treasury bill yield, (ii) Consumer Price Index (CPI), (iii) 1-year prime grade municipal bond yield. The Treasury bill data and the municipal bond data are from Salomon Brothers. The consumer price index is from Citibase. Annualized percentage changes in the CPI are used as proxies for expected inflation.

[6] A technical appendix that provides more detail on each estimator is available from the first author's web site at http://bubba.uta.edu/~crowder.

[7] As pointed out by the referee, Fama (1975) originally suggests this normalization. Fama argues that noise in inflation from the unexpected inflation component is large relative to the noise in the nominal interest rate arising from the stochastic real rate making a regression of inflation on interest rates the more sensible normalization.

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) τ-tests are presented in Table I. Using the critical value for this test of -3.42 given by Dickey and Fuller (1979), the null hypothesis of a unit root cannot be rejected for any of the three series when a lag of 8 or larger is used.[8]

The nonstationarity of the data implies that the use of cointegration techniques is most appropriate in analyzing the Fisher equation. The discussion in Section I suggests that the cointegrating vectors should be of the form $[1 - \beta_{Muni}]'$,where β_{Muni} equals one, for municipal bonds and $[1 - \beta_{TB}]'$, where β_{TB} is greater than one, for Treasury bills. These cointegrating vectors would yield a stationary ex post real interest rate, one with a constant mean. Additionally, if one assumes an average marginal tax rate of 0.3, as implied by the approximate average marginal tax rate computed by Barro and Sahasakul (1986), the cointegrating vector between the 1-year Treasury and the inflation rate should be $[1,-1.42]'$.

These hypotheses are tested using the six different estimators discussed above. Table II presents the estimated Fisher effect parameters for both Treasury bill (T-bill) and municipal bond (Muni) data using each of the six estimators.[9] The distribution of each estimator is asymptotically normal, except for OLS, which is nonstandard. Therefore the standard errors presented are valid for making statistical inference, except for OLS. The OLS standard errors are presented in order to make easy comparisons with other studies.

[8] It is well documented, see Schwert (1987), that postwar U.S. inflation has a large negative moving average (MA) component, see Crowder and Hoffman (1996), Ball and Cecchetti (1990), and Crowder (1997). Schwert (1987) demonstrates that standard unit root tests have extremely large size distortions in these situations. His Monte Carlo results lead to the conclusion that the ADF test with high lag order suggested by Said and Dickey (1984) has the least size distortion.

[9] All of the estimates are produced with constants excluded. Estimation allowing for constants and/or linear trends result in the trends always being statistically insignificant and the constant usually so. The only cases where the constant is significant are in the DOLS1 and OLS1 regressions. The OLS standard errors are usually inappropriate for conventional hypothesis testing in nonstationary regressions and thus can be discounted. The DOLS1 results, however, should be valid, at least asymptotically. It is interesting to note, though, that the estimates of the constant in OLS2 and DOLS2 regressions are insignificant. This may imply that the DOLS1 estimator is not purging the finite sample bias effectively, as suggested by Ng and Perron (1997). Crowder and Hoffman (1995) demonstrate in a Monte Carlo experiment that including a superfluous constant and/or trends in the estimation of the cointegrating vectors may yield severely biased estimates. In any event, allowing constants and/or linear trends in the estimation does not change our qualitative results in any way. For example, the DOLS1 estimate of the municipal bond Fisher effect with constant in regression is 0.343 with a standard error of 0.070. The DOLS1 estimate of the T-bill Fisher effect with a constant included is 0.695 with a standard error of 0.114. These are both much smaller than the estimates obtained without a constant (consistent with the findings in Crowder and Hoffman (1995)) but the relationship between the two Fisher effects is left unchanged, that is, β_{TB} is greater than β_{Muni}.

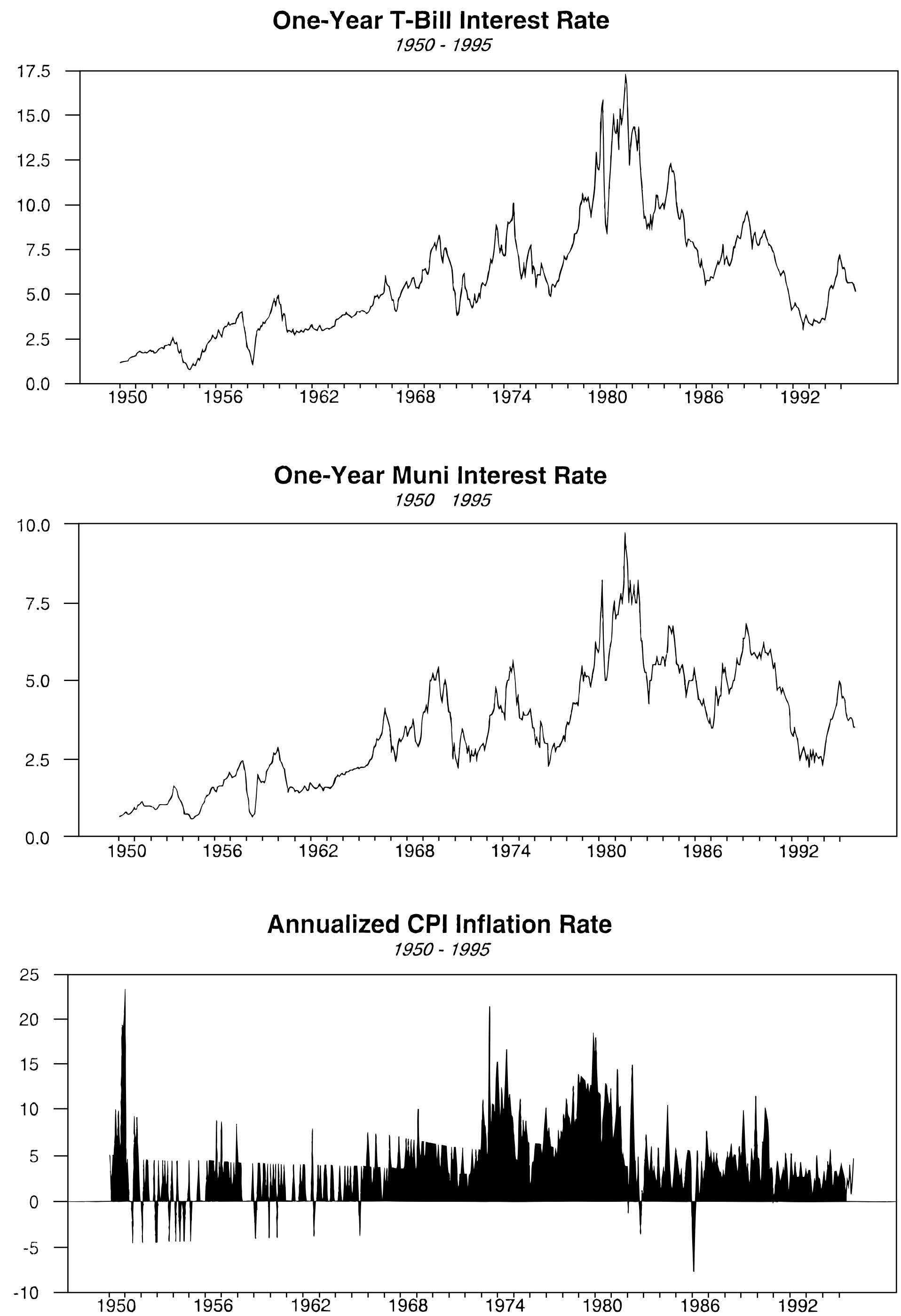

Figure 1. 12-month Treasury bill yield, 12-month municipal bond yield, and 12-month CPI inflation rate.

Table I

Dickey-Fuller τ-Tests

Test statistics are calculated from the regression

$$\Delta X_t = \alpha_0 + \alpha_1 \text{trend} + \alpha_2 X_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta X_{t-j} + \varepsilon_t,$$

where X_t is the T-bill, Muni, or CPI inflation rate, respectively, trend is a linear time trend, and k is the lag length referred to in column 1. For the augmented Dickey-Fuller statistic the 5 percent critical value is -3.42. The sample period is January 1950 to December 1995.

Lag Length	1-Year T-bill	1-Year Muni Bond	1-Year Inflation
4	−2.09	−2.77	−4.95
6	−1.97	−2.68	−3.79
8	−1.92	−2.37	−2.92

The single equation estimators each have two entries in Table II, following the procedure utilized by Fama (1975) and later suggested by Ng and Perron (1997). Those that are labeled with a 1, for example OLS1, are from regressions of interest rates on inflation so that the parameter of interest is the estimated parameter β. The estimators labeled with a 2 are based on regressions with inflation as the regressand and interest rates as the regressor. The parameter of interest is now $1/\beta$.

The OLS estimate of the Fisher effect for the T-bill from the OLS1 specification is quite low and consistent with estimates reported in many earlier studies of the Fisher relationship. The same is true for the Fisher effect estimate for the Muni equation. When the regressor/regressand is reversed (OLS2) the implied Fisher effect from the T-bill equation is 1.471 and that of the Muni equation is 1.009, both of which are remarkably close to their theoretically implied values. Similarly, for DOLS1 the Fisher effect for the T-bill is estimated to be 1.274 and that from the Muni equation is 0.734. Although one cannot reject the null hypothesis that the T-bill Fisher effect is greater than one in this specification, the null that the Muni Fisher effect is equal to one is rejected at less than the 5 percent level of significance. In the DOLS2 specification, the T-bill Fisher effect is estimated to be 1.499. A Wald test of the null hypothesis that this estimate is equal to 1.42, the value suggested by Summers (1983), yields a $\chi^2(1)$ statistic of 0.26 which has a marginal significance level of 61 percent. In the DOLS2 specification of the Muni equation, the estimated Fisher effect is 0.912. The Wald statistic testing the null hypothesis that the implied Fisher effect parameter is equal to one is 0.53 which is again distributed $\chi^2(1)$ with a marginal significance level of 47 percent. Very similar results are found for the ARDL1 versus ARDL2 specifications. The differences between the FM-OLS1 and FM-OLS2 specifications are not as striking.

It is particularly interesting to note that in every case the estimated Fisher effect is smaller for the municipal bond series than for the Treasury bill series. This is consistent with Darby's (1975) extension of the Fisher equa-

Table II

Fisher Effect Estimates

The single equation estimators each have two entries. Those labeled with a 1, for example OLS1, are from regressions of the nominal interest rate, i_t, on inflation, π_{t+1}, so that the parameter of interest is the estimated parameter β. The estimators labeled with a 2 are based on regressions of π_{t+1} on i_t. The parameter of interest is now $1/\beta$ or the inverse of the estimated parameter. Similar notation applies to Dynamic OLS (DOLS), autoregressive distributed lag (ARDL), and fully modified OLS (FM-OLS) estimates. DOLS1 and DOLS2 include eighteen leads and lags with standard errors obtained by using the Newey-West (1987) estimate of the variance based on the Parzen kernel with lag truncation set to 50. The ARDL1 and ARDL2 estimates are calculated from an ARDL(10,6) specification. The FM-OLS1 and FM-OLS2 estimates are obtained using a Var(1) prewhitening of the residuals and an automatic bandwidth selection. The Engle-Yoo (EY) and Johansen maximum likelihood (MLE) estimators include six lags in the vector autoregression (VAR). The sample period is January 1950 to December 1995.

	OLS1	OLS2	DOLS1	DOLS2	ARDL1	ARDL2	FM-OLS1	FM-OLS2	EY	MLE
Panel A: Treasury Bill										
β	0.89	0.68	1.27	0.67	1.33	0.62	1.45	0.61	1.47	1.49
σ_β	0.03	0.02	0.13	0.06	0.21	0.08	0.10	0.13	0.13	0.15
Panel B: Municipal Bond										
β	0.51	0.99	0.73	1.10	0.82	0.99	0.88	0.93	0.89	0.90
σ_β	0.02	0.07	0.09	0.13	0.15	0.16	0.07	0.03	0.09	0.13

tion to account for the effects of taxes on interest income. This result does not depend on a particular estimator. The implication is that fiscal illusion cannot explain the low Fisher effect estimates.

It is also interesting that all but the OLS1 procedure produce estimates of the T-bill Fisher effect that are not statistically different from the values implied by average marginal tax rates in the United States, lending strong support to a full Fisher effect in the United States. Furthermore, all but OLS1 and DOLS1 estimators produce estimates of the Muni Fisher effect that are insignificantly different from the theoretically implied value of one. This evidence strongly suggests that Tobin effects and peso problems may be absent in the long-run relationship between nominal interest rates and inflation.[10]

[10] We also obtain estimates using data that are sampled every January, yielding 46 annual observations. This avoids the overlapping observation problem. Although this problem should not affect our monthly estimates or standard errors for five of the estimators, DOLS, ARDL, EY, FM-OLS and MLE, it may be a problem for OLS. The results using the annual data are qualitatively identical to those using the monthly data. For example, the OLS1 estimate of the T-bill Fisher effect is 1.21, with a standard error of 0.11, and the Muni Fisher effect estimate is 0.69, with a standard error of 0.07. The OLS2 estimate yields an implied Fisher effect of 1.59 for the T-bill and 0.966 for the Muni. Johansen MLE estimates (standard errors in parentheses) are 1.65 (0.20) and 1.03 (0.17), for the T-bill and Muni, respectively. Annual results for all six estimators are available from the authors on request.

III. Conclusion

Empirical estimates of the response of nominal interest rates to changes in inflation, the Fisher effect, have previously been found to be less than the value implied by theory. In this paper we examine whether taxes are important in determining the long-run response of nominal interest rates to changes in (expected) inflation using six different estimators. A finding that the effect of taxes on returns is not incorporated into economic agents' investment calculus would indicate a bizarre violation of rational expectations that Tanzi (1980) calls "fiscal illusion."

The results of this paper provide evidence against fiscal illusion and difference in estimators as possible causes of low Fisher effect estimates. Tobin effects and peso problems are still potential candidates for low Fisher estimates, although we find little evidence of this. However, whatever the cause of low Fisher effects, they impact taxable and tax-exempt bonds in similar ways. In particular, our results demonstrate that, regardless of which estimator of the long-run parameter is employed, the response of tax-free municipal bond rates is always smaller than the response on similar taxable Treasury bills, the difference is statistically significant, and the magnitudes are economically meaningful.

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