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An empirical analysis of term premiums using significance tests for stochastic dominance

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Abstract

The economic significance of term premiums in real returns on US Treasury Bills is examined using recently developed tests for first- and second-order stochastic dominance. The tests place only general restrictions on the preferences of individuals and on the distribution of returns. The results indicate that the two-month real return is preferred to the one-month real return based on both dominance criteria. Other term premiums do not appear to be economically significant.

Keywords: Stochastic Dominance; Term Premium

JEL classification: G10; C10

1. Introduction

The existence of term premiums in returns on US Treasury Bills has long been recognized as an important feature of the term structure (see Roll, 1970, 1971; Fama, 1976a,b, 1984a,b; Shiller et al., 1983; Lauterbach, 1989). It appears that term premiums increase with term-to-maturity and are statistically significant, but the variance of holding-period returns also increases with term-to-maturity. Hence it is not clear that investments in long bills will necessarily be preferred to short bills, even though long bills offer higher returns on average. This paper evaluates the economic significance of real term premiums on US Treasury Bills across corresponding terms-to-maturity, using tests for first- and second-degree stochastic dominance.

Stochastic dominance criteria are conceptually attractive in comparing term premiums, since they allow comparisons between distributions to be made in a very general way. Early empirical studies examining dominance relationships in finance were primarily based on parametric methods, or failed to account for the sampling errors associated with empirical distribution function estimates (Levy and Brooks, 1989). McFadden (1989) has proposed nonparametric dominance test procedures that account

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for sampling errors but these assume that sample returns are i.i.d. and that the samples to be compared are independent. Unfortunately, these assumptions are generally inappropriate for financial data.

This paper applies new distribution-free tests for stochastic dominance that accommodate the contemporaneous and temporal dependence structure that is typically exhibited by returns on financial assets. One general conclusion from previous research that employs dominance criteria is that first-degree stochastic dominance does not provide discriminatory information concerning the relative rankings of assets such as one- to sixth-month Treasury Bills (Levy and Brooks, 1989). However, this conclusion is based on empirical evidence that essentially ignores the positive covariance between returns.

The test procedures used in this paper specify the null hypothesis properly, a feature that is not shared by most other commonly employed procedures (see, for example, Tolley and Pope, 1988; Bishop et al., 1989; McFadden, 1989). Most test procedures make use of the null hypothesis that two distribution or quantile functions are identical. This null hypothesis is not very helpful in providing information about dominance (see Levy, 1992, p. 574). The hypothesis of dominance should be viewed as an hypothesis of inequality in a particular direction between two distributions or quantile functions.

The tests for first- and second-degree stochastic dominance proposed here are based on sample quantile estimates and are a natural extension of Xu et al. (1996) and Xu (1998). Under weak-dependence assumptions, these estimates have a joint asymptotic normal distribution with a complicated variance–covariance structure, consistent estimates of which can be obtained using the moving-block bootstrap (MBB) developed by Kunsch (1989) and Liu and Singh (1992). Tests for first- and second-degree stochastic dominance are then based on a pseudo-likelihood-ratio test procedure for joint inequality restrictions developed by Kodde and Palm (1986).

Section 2 reviews stochastic dominance concepts and develops the tests. Section 3 presents the empirical results. Finally, some concluding remarks are offered in Section 4.

2. Stochastic dominance tests

Two classes of utility functionals are U_1 and U_2. U_1 includes all the utility functions u for which $u' \geq 0$; U_2 includes all functions u for which $u' \geq 0$ and $u'' \leq 0$.[1] Let the support of each distribution function be $[a,b]$, where $a < b$ are finite real numbers. Let F_X and F_Y be the marginal distribution functions of random returns X and Y, respectively. Two stochastic dominance criteria are:

$$X \text{ dominates } Y \text{ in the first-degree } (XD_1Y) \text{ if } F_Y(w) - F_X(w) \geq 0 \quad \forall w \in [a,b],$$

$$X \text{ dominates } Y \text{ in the second-degree } (XD_2Y) \text{ if } \int_a^w [F_Y(t) - F_X(t)]\mathrm{d}t \geq 0 \quad \forall w \in [a,b].$$

The inequality in each of the definitions is taken to hold at least once

$$XD_1Y \Leftrightarrow E_{F_X}u(X) \geq E_{F_Y}u(Y) \quad \forall u \in U_1 \text{ and } XD_2Y \Leftrightarrow E_{F_X}u(X) \geq E_{F_Y}u(Y) \quad \forall u \in U_2.$$

[1]See Hadar and Russell (1969), Hanoch and Levy (1969), and Rothschild and Stiglitz (1970).

The dominance criteria can be equivalently expressed in terms of quantile functions.[2] Let Q_X and Q_Y be the quantile functions of random returns X and Y. XD_1Y if $Q_X(p) - Q_Y(p) \geq 0\ \forall p \in [0,1]$; XD_2Y if $\int_0^p [Q_X(t) - Q_Y(t)]\mathrm{d}t \geq 0\ \forall p \in [0,1]$. Alternatively, XD_2Y if $\Psi_X(p) - \Psi_Y(p) \geq 0\ \forall p \in [0,1]$, where $\int_0^p Q_X(t)\mathrm{d}t = \Psi_X(p)$ and $\int_0^p Q_Y(t)\mathrm{d}t = \Psi_Y(p)$.

K corresponding points on each of the estimated sample quantile functions are denoted by the $K \times 1$ vectors $\hat{Q}_X$ and $\hat{Q}_Y$. These are assembled in the $2K \times 1$ vector $\hat{Q}_Z = [\hat{Q}'_X \vdots \hat{Q}'_Y]'$. Under sufficient regularity, the asymptotic distribution of $\sqrt{T}Q_Z$ is $N(\xi,\Lambda)$ whereupon if

$$(\hat{Q}_X - \hat{Q}_Y) = H\hat{Q}_z,\ \ H = [I_K \vdots -I_K],$$

then

$$\sqrt{T}H\hat{Q}_Z \xrightarrow{d} N(H\xi, H\Lambda H')$$

(see Sen, 1972). The elements of Λ depend on the serial correlation structure of the data and are, in principle, difficult to estimate. The moving-block bootstrap (MBB) proposed by Kunsch (1989) and Liu and Singh (1992) is used to estimate Λ consistently.

When X and Y represent a pair of holding-period returns, XD_1Y is tested on the null hypothesis H_0: $Q_X(p) - Q_Y(p) \geq 0\ \forall p \in [0,1]$ against the alternative hypothesis H_a: $Q_X(p) - Q_Y(p) \ngeq 0$ for at least one p. Let $\Delta = [(\hat{Q}_X - \hat{Q}_Y) - (\tilde{Q}_X - \tilde{Q}_Y)]$ where the circumflex denotes unrestricted estimates and the tilde represents restricted estimates. Restricted estimates are obtained by minimizing

$$[(\hat{Q}_X - \hat{Q}_Y) - (Q_X - Q_Y)]'\left[\frac{1}{T}H\hat{\Lambda}H'\right]^{-1}[(\hat{Q}_X - \hat{Q}_Y) - (Q_X - Q_Y)],$$

$$\text{s.t.}\ \ Q_X - Q_Y \geq 0.$$

Then the test-statistic $c_1 = \Delta'[(1/T)H\hat{\Lambda}H']^{-1}\Delta$ is asymptotically distributed as a weighted sum of χ^2-variates with different degrees of freedom; see Kodde and Palm (1986).

A lower bound q_1 for the critical value of c_1 is obtained by choosing a significance level α and setting degrees of freedom (df) equal to one [i.e., $\alpha = 1/2\,Pr(\chi^2_K \geq q_1)$]. An upper bound q_u for the critical value is obtained by choosing a significance level α and setting df equal to K [i.e., $\alpha = 1/2\,Pr(\chi^2_{K-1} \geq q_u) + 1/2\,Pr(\chi^2_K \geq q_u)$]. Decision rules for the statistic c_1 are: if c_1 exceeds the upper bound value q_u, reject H_0; and if c_1 is smaller than the lower bound value q_l, do not reject H_0. If c_1 is in the inconclusive region, Monte Carlo simulations should be used to compute the weights for the case $K \geq 8$; see Wolak (1989a), (1989b).

XD_2Y is tested based on H_0: $\Psi_X(p) - \Psi_Y(p) \geq 0\ \forall p \in [0,1]$ against H_α: $\Psi_X(p) - \Psi_Y(p) \ngeq 0$ for at least one p. Before describing the XD_2Y test-statistic, it is useful to define a cumulative quantile generator B, a $(K \times K)$ lower triangular matrix

$$B = \begin{bmatrix} 1 & 0 & 0 & \cdots & 0 & 0 \\ 1 & 1 & 0 & \cdots & 0 & 0 \\ \vdots & \vdots & \vdots & \ddots & \vdots & \vdots \\ 1 & 1 & 1 & \cdots & 1 & 0 \\ 1 & 1 & 1 & \cdots & 1 & 1 \end{bmatrix}$$

[2] Levy and Kroll (1978) demonstrate the equivalence between the distribution and the quantile formulations.

Given the number of chosen quantile points K, B premultiples $\hat{Q}$ yielding a K-variate vector of cumulative sample quantiles

$$\begin{aligned} B\hat{Q} &= \left[\hat{Q}(p_1), \sum_{i=1}^{2} \hat{Q}(p_i), \ldots, \sum_{i=1}^{K} \hat{Q}(p_i) \right]' \\ &= [\hat{\Psi}(p_1), \hat{\Psi}(p_2), \ldots, \hat{\Psi}(p_K)]' \\ &= \hat{\Psi}, \end{aligned}$$

where $p_{i+1} - p_i = p_{j+1} - p_j$ for all $i,j = 0,1,\ldots,K-1$, $p_0 = 0$, $p_K = 1$. The variance–covariance matrix of $B(\hat{Q}_x - \hat{Q}_Y)$ is $1/T\ BH\hat{\Lambda}H'B'$. The test-statistic for XD_2Y is given by $c_2 = (B\Delta)'[1/T\ BH\hat{\Lambda}H'B']^{-1}(B\Delta)$. The restricted estimates in $B\Delta$ are calculated by minimizing

$$B[(\hat{Q}_X - \hat{Q}_Y) - (Q_X - Q_Y)]' \left[\frac{1}{T} BH\hat{\Lambda}H'B' \right]^{-1} B[(\hat{Q}_X - \hat{Q}_Y) - (Q_X - Q_Y)],$$

$$\text{s.t. } B(Q_X - Q_Y) \geq 0.$$

The test-statistic c_2 is asymptotically distributed as the weighted sum of χ^2-variates. The decision rules based on c_2 are the same as those for the test-statistic c_1.

Practical application of these tests requires a convenient estimate of the variance–covariance matrix $\hat{\Lambda}$ for $(\hat{Q}_X - \hat{Q}_Y)$; for this the MBB is applied. Consistency of the MBB is achieved under the regularity conditions, if the number of observations in each block, b, approaches infinity with T in such a way that the number of moving blocks, $k = [T/b]$, also approaches infinity with T. In practice, randomly drawn blocks of b adjacent observations are used to form a resample for computing one $(\hat{Q}_X - \hat{Q}_Y)$. When this is repeated many times, the collection of $(\hat{Q}_X - \hat{Q}_Y)$ calculations can then be used to compute the variance–covariance matrix $H\hat{\Lambda}H'$, and hence $BH\hat{\Lambda}H'B'$.

3. Empirical results

The yields of 1–6 month Treasury Bills for the period of 1952:02–1987:02 were obtained from McCulloch.[3] To transform yields to nominal holding-period returns, the equation of Shiller (1990) has been used. Dickey–Fuller test results suggest that nominal holding-period returns have a unit root. Nominal holding-period returns are transformed to real holding-period returns ($h^r(i)$, $i = 1,2,\ldots,6$) which do not appear to have a unit root. Table 1 indicates that both the mean and variance of real holding-period returns tend to increase as the term-to-maturity becomes longer. Estimated skewness and kurtosis indicate that the return distributions are not normal and that a distribution-free method is warranted. Table 2 indicates that real holding-period returns have a non-trivial serial correlation structure, and that the extra complexity associated with the MBB variance–covariance matrix estimate is warranted. Table 3 shows that the values of the sample correlation coefficients between two real

[3]The data are printed in the appendix of Shiller (1990).

Table 1
Basic statistics of real returns: 1954:02–1987:01

Series	Mean	Variance	Skewness	Kurtosis
$h^r(1)$	0.06780	0.08100	0.06921	1.90701
$h^r(2)$	0.11238	0.08584	0.14013	1.78988
$h^r(3)$	0.13600	0.10052	0.07462	2.00694
$h^r(4)$	0.15140	0.11646	0.32890	1.86587
$h^r(5)$	0.16031	0.13935	0.50792	2.58086
$h^r(6)$	0.17006	0.16877	0.75224	3.68112

Note: $h^r(i)$, $i = 1,2,\ldots,6$, represent the real holding-period returns of i-month US Treasury Bills.

Table 2
Autocorrelation functions for real returns: 1954:02–1987:01

	Lag 5	Lag 10	Lag 15	Lag 20	Lag 25	Lag 30	Lag 35	Lag 40	Lag 45	Lag 50
$h^r(1)$	0.313	0.353	0.281	0.236	0.089	0.116	0.042	0.058	−0.073	−0.051
$h^r(2)$	0.323	0.338	0.314	0.241	0.124	0.022	0.076	0.061	−0.062	−0.041
$h^r(3)$	0.326	0.321	0.317	0.208	0.119	0.090	0.074	0.016	−0.048	−0.024
$h^r(4)$	0.322	0.273	0.256	0.147	0.100	−0.032	0.056	0.045	−0.045	−0.048
$h^r(5)$	0.306	0.248	0.220	0.098	0.062	0.049	0.022	−0.007	−0.049	−0.036
$h^r(6)$	0.273	0.211	0.171	0.053	0.041	−0.055	0.014	0.018	−0.045	0.023

Note: $h^r(i)$, $i = 1,2,\ldots,6$, represent the real holding-period returns of i-month US Treasury Bills.

holding-period returns range from 0.70827 to 0.98151. In addition, the correlation decreases as the difference in term-to-maturity increases.

The economic significance of term premiums is investigated using tests for stochastic dominance among selected pairs of the holding-period returns. The MBB estimates of the variance–covariance matrix for the difference between two sets of quantile estimates are computed with 200 replications and various block sizes ($b=30$, 40, 50, and 60). The test results are then computed using the MBB variance–covariance matrices. For each pair, dominance relations in both directions are tested. The reported test-statistics are calculated on the basis of 20 equally-spaced quantile estimates. To test the null hypotheses of first- or second-degree stochastic dominance at $\alpha=0.05$, each null hypothesis will be rejected if the test-statistic is greater than 30.841; it will not be rejected if the test-statistic is less

Table 3
Correlation matrix for real returns: 1954:02–1987:01

	$h^r(1)$	$h^r(2)$	$h^r(3)$	$h^r(4)$	$h^r(5)$	$h^r(6)$
$h^r(1)$	1.00000					
$h^r(2)$	0.96835	1.00000				
$h^r(3)$	0.90843	0.97070	1.00000			
$h^r(4)$	0.84126	0.91277	0.96482	1.00000		
$h^r(5)$	0.77704	0.85711	0.93046	0.98151	1.00000	
$h^r(6)$	0.70827	0.79990	0.88763	0.95484	0.99064	1.00000

Note: $h^r(i)$, $i = 1,2,\ldots,6$, represent the real holding-period returns of i-month US Treasury Bills.

Table 4
Test-statistics of first-degree stochastic dominance c_1 with different MBB block sizes

H_0	$b=30$	$b=40$	$b=50$	$b=60$
$h^r(2)D_1h^r(1)$	0.000	0.000	0.000	0.000
$h^r(3)D_1h^r(1)$	0.051	0.049	0.054	0.064
$h^r(4)D_1h^r(1)$	0.000	0.000	0.000	0.000
$h^r(5)D_1h^r(1)$	0.004	0.004	0.005	0.004
$h^r(6)D_1h^r(1)$	0.107	0.107	0.118	0.119
$h^r(1)D_1h^r(2)$	12.564	11.069	10.726	10.930
$h^r(1)D_1h^r(3)$	4.370	5.178	6.133	4.802
$h^r(1)D_1h^r(4)$	4.761	4.583	5.181	4.992
$h^r(1)D_1h^r(5)$	3.150	3.715	3.405	3.826
$h^r(1)D_1h^r(6)$	2.268	2.426	2.361	2.419

Note: $h^r(i)D_1h^r(j)$ denotes that the i-month real holding-period return dominates, in the first-degree, the j-month real holding period-return. The number of points selected, K, is 20. The number of resamples in the moving block bootstrap is 200. At $\alpha=0.05$, H_0, under which either $Q_Y(p)-Q_X(p)\geq 0\ \forall p\epsilon[0,1]$ or $Q_X(p)-Q_Y(p)\geq 0\ \forall p\epsilon[0,1]$, will be rejected if the test-statistic is greater than 30.841; will not be rejected if the test-statistic is less than 2.706.

than 2.706. If the test-statistic is in between these two critical values and there is a need for a clearer decision rule, the weights in the test-statistic must then be calculated.

Table 4 reports dominance relations under the null hypothesis and test-statistics corresponding to $b=30$, 40, 50, and 60. The test-statistics show that the longer-term returns $h^r(2)$, $h^r(3)$, $h^r(4)$, $h^r(5)$, and $h^r(6)$ dominate, in the first-degree (FSD), the one-month return $h^r(1)$, at the 5% significance level. Table 5 shows the test results for second-degree stochastic dominance (SSD) relations among selected pairs of holding-period returns. These test-statistics show that the longer-term returns $h^r(2)$, $h^r(3)$, $h^r(4)$, $h^r(5)$, and $h^r(6)$ generally dominate, in the second-degree, the one-month return $h^r(1)$. This is consistent with the fact that FSD implies SSD. All of the null hypotheses cannot be rejected at the 5% significance level.

Table 5
Test-statistics of second-degree stochastic dominance c_2 with different MBB block sizes

H_0	$b=30$	$b=40$	$b=50$	$b=60$
$h^r(2)D_2h^r(1)$	0.000	0.000	0.000	0.000
$h^r(3)D_2h^r(1)$	0.054	0.051	0.057	0.057
$h^r(4)D_2h^r(1)$	0.000	0.000	0.000	0.000
$h^r(5)D_2h^r(1)$	0.000	0.000	0.000	0.000
$h^r(6)D_2h^r(1)$	0.109	0.191	0.126	0.126
$h^r(1)D_2h^r(2)$	10.087	8.719	6.360	7.707
$h^r(1)D_2h^r(3)$	3.161	2.286	2.742	2.476
$h^r(1)D_2h^r(4)$	2.888	2.992	2.727	2.788
$h^r(1)D_2h^r(5)$	1.381	1.584	1.855	1.494
$h^r(1)D_2h^r(6)$	0.752	0.818	0.756	0.792

Note: $h^r(i)D_2h^r(j)$ denotes that the i-month real holding-period return dominates, in the second-degree, the j-month real holding-period return. The number of points selected, K, is 20. The number of resamples in the moving block bootstrap is 200. At $\alpha=0.05$, H_0, under which either $\Psi_Y(p)-\Psi_X(p)\geq 0\ \forall p\epsilon[0,1]$ or $\Psi_X(p)-\Psi_Y(p)\geq 0\ \forall p\epsilon[0,1]$, will be rejected if the test-statistic is greater than 30.841; will not be rejected if the test-statistic is less than 2.706.

Table 6
p-Values of test-statistics

H_0	$b=30$	$b=40$
$h^r(1)D_1h^r(2)$	**0.023**	**0.041**
$h^r(1)D_1h^r(3)$	0.235	0.196
$h^r(1)D_1h^r(4)$	0.222	0.238
$h^r(1)D_1h^r(5)$	0.341	0.352
$h^r(1)D_2h^r(2)$	**0.055**	**0.098**
$h^r(1)D_2h^r(3)$	0.391	0.524
$h^r(1)D_2h^r(4)$	0.426	0.421

Note: $h^r(i)D_sh^r(j)$ denotes that the i-month real holding-period return dominates, in the s-degree, the j-month real holding-period return. If the p-value is less than the chosen significance level, say 10%, then the null hypothesis is rejected. The bold numbers are less than the chosen significance level.

These results may be interpreted as evidence that the longer-term returns do indeed dominate the shorter returns; but the results may also arise as a consequence of lack of power. Accordingly, we consider the tests in the opposite direction. Tables 4 and 5 reveal that there is less numerical evidence that the opposite dominance relations hold because the test-statistics are generally higher than the lower bound, and lower than the upper bound, of the critical value. Table 6 reports the p-values associated with the exact tests for $b=30$, 40. It turns out that only the two-month return dominates the one-month return in the first- and second-degree, i.e. the null hypothesis cannot be rejected at the 10% significance level. This means that, relative to the two-month return, the one-month return does not embody sufficient risk premium.

In comparison with Levy and Brooks (1989), it is interesting to note that at least one longer-term bill dominates the shortest bill in both first- and second-degree. Levy and Brooks ignored the sampling errors associated with the distribution function estimates, yet concluded that no assets are dominated in the first-degree for their entire sample period. Our results indicate that the shortest bill is weakly dominated by the two-month bill, when sampling errors are taken into account.

4. Concluding remarks

In this paper, new distribution-free tests for first- and second-degree stochastic dominance have been described and applied to evaluating the economic significance of real term premiums. These test procedures are advantageous because (i) they permit the returns of different assets to be dependent, (ii) they do not require sample observations for a particular return to be i.i.d., and (iii) the null hypothesis is correctly specified.

The tests have been applied to McCulloch's US Treasury Bill data as given in Shiller (1990). The results for first- and second-degree stochastic dominance suggest that only the holding-period return of the one-month Treasury Bill is significantly dominated, in both first- and second-degree, by the holding-period return on the second shortest term-to-maturity in the data set. As a result, the two-month bill is strictly preferred to the one-month bill, in both first- and second-degree in terms of holding-period returns. Overall, the tests provide no clear evidence concerning dominance relationships among the other maturities. The results have been contrasted with those of Levy and Brooks

(1989) and illustrate a useful application of dominance testing when the samples are associated and each sample contains temporally dependent observations. The Levy–Brooks results do not allow for sampling fluctuations, while the results in this paper do; but the results of this paper also accommodate the association and dependence that are characteristic of financial data.

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References

Bishop, J.A., Chakraborti, S., Thistle, P.D., 1989. Statistical inference, income distributions, and social welfare. Research on Income Inequality 1, 49–82.

Fama, E., 1976a. Inflation uncertainty and expected returns on treasury bills. Journal of Political Economy 84, 427–448.

Fama, E., 1976b. Forward rates as predictors of future spot rates. Journal of Financial Economics 3, 361–377.

Fama, E., 1984a. The information in the term structure. Journal of Financial Economics 13, 509–528.

Fama, E., 1984b. Term premium in bond returns. Journal of Financial Economics 13, 529–546.

Hadar, J., Russell, W.R., 1969. Rules for ordering uncertain prospects. American Economic Review 59, 25–34.

Hanoch, G., Levy, H., 1969. The efficiency analysis of choices involving risk. Review of Economic Studies 36, 335–346.

Kodde, D.A., Palm, F.C., 1986. Wald criteria for jointly testing equality and inequality restrictions. Econometrica 50, 1243–1248.

Kunsch, H., 1989. The jackknife and the bootstrap for general stationary observations. Annals of Statistics 17, 1217–1241.

Lauterbach, B., 1989. Consumption volatility, production volatility, spot-rate volatility, and returns on treasury bills and bonds. Journal of Financial Economics 24, 155–179.

Levy, H., 1992. Stochastic dominance and expected utility. Management Science 38, 555–593.

Levy, H., Brooks, R., 1989. An empirical analysis of term premiums using stochastic dominance. Journal of Banking and Finance 13, 245–260.

Levy, H., Kroll, Y., 1978. Ordering uncertain options with borrowing and lending. Journal of Finance 33, 553–574.

Liu, R.Y., Singh, K., 1992. Moving blocks jackknife and bootstrap capture weak dependence. In: LePage, R., Billard, L. (Eds.), Exploring the Limits of Bootstrap. Wiley, New York, pp. 225–248.

McFadden, D., 1989. Testing for stochastic dominance. In: Fomby, T.B., Seo, T.K. (Eds.), Studies in the Economics of Uncertainty: In Honor of Josef Hadar. Springer, New York, pp. 113–134.

Roll, R., 1970. The Behavior of Interest Rates: The Application of the Efficient Market Model to U.S. Treasury Bills. Basic Books.

Roll, R., 1971. Investment diversification and bond maturity. Journal of Economic Theory 2, 225–243.

Rothschild, M., Stiglitz, J.E., 1970. Increasing risk: I. A definition. Journal of Economic Theory 2, 225–243.

Sen, P.K., 1972. On the Bahadur representation of sample quantiles for sequences of ϕ-mixing random variables. Journal of Multivariate Analysis 2, 77–95.

Shiller, R.J., 1990. The term structure of interest rates. In: Friedman, B.M., Han, F.H. (Eds.), Handbook of Monetary Economics, Vol. I. Chapt. 13, pp. 627—722.

Shiller, R.J., Campbell, J.Y., Schoenholtz, K.L., 1983. Forward rates and future policy: interpreting the term structure of interest rates. Brookings Papers on Economic Activity 1, 173–217.

Tolley, H.D., Pope, R.D., 1988. Testing for stochastic dominance. American Journal of Agriculture Economics 70, 693–700.

Wolak, F.A., 1989a. Local and global testing of linear and nonlinear inequality constraints in nonlinear econometric models. Econometric Theory 5, 1–35.

Wolak, F.A., 1989b. Testing inequality constraints in linear econometric models. Journal of Econometrics 41, 205–235.

Xu, K., Fisher, G., Willson, D., 1996. Testing first- and second-order stochastic dominance. Canadian Journal of Economics 29, s526–s564.

Xu, K., 1998. Asymptotic distribution-free statistical test for generalised Lorenz curves: an alternative approach. Journal of Income Distribution (in preparation).