Testing for mean-variance spanning
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Published in:
Journal of Empirical Finance

Document version:
Peer reviewed version

Publication date:
2001

Link to publication

Citation for published version (APA):

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Abstract

In this paper we present a survey on the various approaches that can be used to test whether the mean-variance frontier of a set of assets spans or intersects the frontier of a larger set of assets. We analyze the restrictions on the return distribution that are needed to have mean-variance spanning or intersection. The paper explores the duality between mean-variance frontiers and volatility bounds, analyzes regression based test procedures for spanning and intersection, and shows how these regression based tests are related to tests for mean-variance efficiency, performance measurement, optimal portfolio choice, and specification error bounds.

JEL: G110, G120
Keywords: mean-variance spanning, portfolio choice, volatility bounds, performance measurement
1 Introduction

In recent years the finance literature has witnessed an increasing use of tests for mean-variance spanning and intersection, as introduced by Huberman and Kandel (1987). In this paper we will provide a survey of the literature on testing for mean-variance spanning and intersection, as well as of its relationships with volatility bounds, tests for mean-variance efficiency, performance evaluation, and the specification error bounds that have recently been proposed by Hansen and Jagannathan (1997). There exists a vast literature on most of these subjects and the intention here is not to give a complete overview, but merely to illustrate that the concept of mean-variance spanning and intersection provides a framework in which many other results can be understood.

The literature on mean-variance spanning and intersection analyzes the effect that the introduction of additional assets has on the mean-variance frontier. If the mean-variance frontier of the benchmark assets and the frontier of the benchmark plus the new assets have exactly one point in common, this is known as intersection. This means that there is one mean-variance utility function for which there is no benefit from adding the new assets. If the mean-variance frontier of the benchmark assets plus the new assets coincides with the frontier of the benchmark assets only, there is spanning. In this case no mean-variance investor can benefit from adding the new assets to his (optimal) portfolio of the benchmark assets only. For instance, De Santis (1995) and Cumby and Glen (1990) consider the question whether US-investors can benefit from international diversification. Taking the viewpoint of a US-investor who initially only invests in the US, these authors study the question whether they can enhance the mean-variance characteristics of their portfolio by also investing in other (developed) markets. Similarly, taking the perspective of a US-investor who invests in the US and (possibly) in other developed markets such as Japan and Europe, De Santis (1994), Bekaert and Urias (1996), Errunza, Hogan and Hung (1998), and DeRoon, Nijman and Werker (2001) e.g., investigate whether the investors can improve upon their mean-variance portfolio by investing in emerging markets. As a nal example, Glen and Jorion (1993) investigate whether mean-variance investors with a well-diversi ed international portfolio of stocks and bonds should add currency futures to their portfolio, i.e., whether or not they should hedge the currency risk that arises from their positions in stocks and bonds.

As shown by De Santis (1994), Ferson, Foerster, and Keim (1993), Ferson
(1995) and Bekaert and Urias (1996), the hypothesis of mean-variance spanning and intersection can be reformulated in terms of the volatility bounds introduced by Hansen and Jagannathan (1991). In that case, the interest is in the question whether a set of additional assets contains information about the volatility of the pricing kernel or the stochastic discount factor that is not already present in the initial set of assets considered by the econometrician. For instance, in the case of emerging markets, the question is whether considering returns from the US-market together with returns from emerging markets produces tighter volatility bounds on the stochastic discount factor than returns from the US-market only.

The duality between mean-variance frontiers and volatility bounds for the stochastic discount factors will be the subject of Section 2. The analysis provided in that section will allow us to study mean-variance spanning and intersection, both in terms of mean-variance frontiers and in terms of volatility bounds. The concept of mean-variance spanning and intersection will formally be introduced in Section 3. In that section it will be also be shown how simple regression techniques can be used to test for mean-variance spanning and intersection. In Section 4 we will consider how conditioning information can be incorporated in the test procedures. In Section 5 we will show how deviations from mean-variance intersection and spanning can be interpreted in terms of performance measures like Jensen's alpha and the Sharpe ratio, and how the regression tests for intersection can be used to derive the new optimal portfolio weights. In Section 6 we provide a brief discussion of the specification error bound introduced by Hansen and Jagannathan (1997) and how this is related to mean-variance intersection. As with the performance measures in Section 5, specification error bounds are especially of interest when there is no intersection. This paper will end with a summary.

2 Volatility bounds and the duality with mean-variance frontiers

The purpose of this section is to introduce volatility bounds and mean-variance frontiers and to show the duality between these two frontiers. Because mean-variance spanning and intersection can be defined from volatility bounds as well as from mean-variance frontiers, this section provides a basis
for the analysis of mean-variance spanning and intersection in the remainder of the paper.

2.1 Volatility bounds

Suppose an investor chooses his portfolio from a set of $K$ assets, with current prices given by the $K$-dimensional vector $P_t$ and with payoffs in the next period given by the vector $P_{t+1}$ (including dividends and the like). Returns $R_{i,t+1}$ are payoffs with prices equal to one. Assuming there are no market frictions such as short sale constraints and transaction costs and assuming that the law of one price holds, there exists a stochastic discount factor or pricing kernel, $M_{t+1}$, such that

$$E[M_{t+1}R_{t+1} | I_t] = \Pi_K;$$

(1)

where $\Pi_K$ is a $K$-dimensional vector containing ones, and $I_t$ is the information set that is known to the investor at time $t$. In the sequel we will use $E[. | I_t]$ as shorthand notation for $E[. | I_t]$.

Apart from the law of one price, an alternative way to motivate (1) is to look at the discrete time consumption and portfolio problem that an investor solves:

$$\max_{w_t, C_t} \mathbb{E}_t \left[ \sum_{j=0}^{\infty} \beta_j U(C_{t+j}) \right];$$

(2)

s.t. $W_{t+1} = w_t^0 R_{t+1} (W_t | C_t);$

$w_t^0 \Pi_K = 1; \beta_t$

where $C_t$ is consumption at time $t$, $W_t$ is the wealth owned by the investor at time $t$, $\beta_t$ is the subjective discount factor of the investor, and $w_t$ is the $K$-dimensional vector of portfolio weights that the investor chooses. The function $U(C_t; C_{t+1}; \ldots) = \sum_{j=0}^{\infty} \beta_j U(C_{t+j})$ is a strictly increasing and concave time-separable utility function. The first order conditions of problem (2) imply that

$$M_{t+1} = \left( \frac{U_q(C_{t+1})}{U_q(C_t)} \right)^{C_t^{opt}, w_t^{opt}};$$

is a valid stochastic discount factor with $U_q(.)$ being the first derivative of $U$. Thus, one way to think about the stochastic discount factor or pricing

\footnote{Replacing the law of one price with the stronger condition that there are no arbitrage opportunities, we would also have that $M_{t+1} > 0$.}
kernel is as the intertemporal marginal rate of substitution (IMRS). This interpretation of the pricing kernel is more restrictive than the law of one price though, since it also implies that \( M_{t+1} > 0 \).

In many of the problems we consider in this paper, it is convenient to look at a more simple portfolio problem. Usually we will restrict ourselves to one-period portfolio problems, where the agent maximizes his indirect utility of wealth function (see, e.g., Ingersoll (1987), p.66):

\[
\max_{f \in \mathcal{W}} E_t[u(W_{t+1})];
\]

s.t. \( W_{t+1} = W_t w^R_{t+1}; \)

\( w^\mathcal{K} = 1; \)

In this case a valid stochastic discount factor is \( W_t \mathcal{E} u^0(W_{t+1}) = \mathcal{E}, \) with \( u^0(\cdot) \) being the first derivative of the indirect utility function evaluated at the optimal portfolio choice, and \( \mathcal{E} \) the Lagrange multiplier for the restriction that \( w^\mathcal{K} = 1. \)

The expectation of the stochastic discount factor will be denoted by \( \mathcal{E} \), i.e., \( \mathcal{E} = E_t[M_{t+1}]. \) The name stochastic discount factor refers to the fact that \( M_{t+1} \) discounts payoffs differently in different states of the world. To illustrate this, using the definition of covariance, (1) can be rewritten as

\[
\mathcal{E} = E_t[M_{t+1}R_{t+1}] = \mathcal{E} E_t[R_{t+1}] + \text{Cov}_t[R_{t+1}; M_{t+1}]; \quad (3)
\]

The first term in (3) uses \( \mathcal{E} \) to discount the expected future payoffs, while the second term is a risk adjustment (recall that \( \mathcal{E} \) is the price-vector of the returns \( R_{t+1} \)). Accordingly, risk premia are determined by the covariance of asset payoffs with \( M_{t+1} \). If one of the assets is a risk free asset with return \( R^f_t \), then it follows from the conditional expectation in (1) that \( R^f_t = 1 = \mathcal{E} \).

In the sequel we will usually not impose the presence of such a risk free asset. If a risk free asset is available however, then we can always substitute \( 1 = R^f_t \) for \( \mathcal{E} \).

Equation (1) is the starting point for most asset pricing models. In fact, differences in asset pricing models can be interpreted as differences in the function that each model assigns to \( M_{t+1} \) (see, e.g., Cochrane (1997)). Since each valid stochastic discount factor has to satisfy (1), observed asset returns can be used to derive information about these discount factors. For instance, following Hansen and Jagannathan (1991) it is possible to derive a lower bound on the variance of \( M_{t+1} \), that each valid stochastic discount factor
has to satisfy, which is known as the volatility bound. To see this, we start from the unconditional version of (1), and leave out the time subscripts for the expectations and (co)variance operators, as well as for \( v \). In this paper, the expectation of the stochastic discount factor will usually be a free parameter. We will denote all discount factors that satisfy (1) and that have unconditional expectation \( v \) with \( M(v)_{t+1} \), and derive a lower bound for the variance of each \( M(v)_{t+1} \).

Let the unconditional expectation and covariance matrix of the returns \( R_{t+1} \) be given by \( \bar{1}_R \) and \( \Omega_{RR} \) respectively, and assume that all returns are independently and identically distributed (i.i.d.), so that the expectations and covariances do not vary over time. This assumption will be relaxed in Section 4 of this paper. Given the set of asset returns \( R_{t+1} \), let \( m_R(v)_{t+1} \) be a candidate stochastic discount factor that has expectation \( v \) and that is linear in the asset returns:

\[
m_R(v)_{t+1} = v + ' (v)^{1}_R R_{t+1} \quad (4)
\]

where we write \( ' \) (v) to indicate that these coefficients are a function of the expectation of \( M(v)_{t+1} \). Substituting (4) into (1), we obtain:

\[
' (v) = \Omega_{R}^{-1} \bar{1}_R \Omega_{RR}^{-1} (\Omega_{R}^{1}_R R_{t+1} - v) \quad (5)
\]

Since both \( M(v)_{t+1} \) and \( m_R(v)_{t+1} \) satisfy (1) we have that

\[
E[(M(v)_{t+1} \mid m_R(v)_{t+1}) R_{t+1}] = 0
\]

so the difference between any \( M(v)_{t+1} \) that satisfies (1) and \( m_R(v)_{t+1} \) is orthogonal to \( R_{t+1} \) and therefore to \( m_R(v)_{t+1} \) itself. This implies for the variance of \( M(v)_{t+1} \) that:

\[
\text{Var}[M(v)_{t+1}] = \text{Var}[m_R(v)_{t+1}] + \text{Var}[(M(v)_{t+1} \mid m_R(v)_{t+1})]
\]

(6)

which shows that \( m_R(v)_{t+1} \) has the lowest variance of all valid stochastic discount factors \( M(v)_{t+1} \). This minimum variance can be obtained by combining (4) and (5):

\[
\text{Var}[m_R(v)_{t+1}] = (\Omega_{R}^{-1} v \Omega_{R}^{1}_R) \Omega_{RR}^{-1} (\Omega_{R}^{1}_R v \Omega_{R}^{1})
\]

(7)

Thus, any pricing model that aims to price the assets \( R_{t+1} \) correctly, has to yield a pricing kernel that, for a given \( v \), has a variance at least as large as (7). Equivalently, if we know that agents choose their optimal portfolio from
the assets that are in $R_{t+1}$, then (7) gives the minimum amount of variation of their IMRS that is needed to be consistent with the distribution of asset returns. Luttmer (1996) extends this kind of analysis taking into account market frictions such as short sales constraints and transaction costs. For the frictionless markets setting, Snow (1991) provides a similar analysis to derive bounds on other moments of the discount factor as well, and Bansal and Lehmann (1997) provide a bound on the mean of the logarithm of the pricing kernel, using growth optimal portfolios. Balduzzi and Kallal (1997) show how additional knowledge about risk premia may lead to sharper bounds on the volatility of the discount factor and Balduzzi and Robotti (2000) use the minimum variance discount factor to estimate risk premia associated with economic risk variables. Finally, Bekaert and Liu (1999) and Ferson and Siegel (1997) study the use of conditioning information to derive optimally scaled volatility bounds.

2.2 Duality between volatility bounds and mean-variance frontiers

In the previous section we derived the minimum amount of variation in stochastic discount factors that is needed to be consistent with the distribution of asset returns. In this section we will show that there is a close correspondence between these volatility bounds and mean-variance frontiers and that stochastic discount factors that correspond to mean-variance optimizing behavior are the stochastic discount factors with the lowest volatility. Mean-variance optimizing behavior is a special case of the portfolio problem considered before, where the problem the agent faces is $\max_{f,w} E[u(W_{t+1})]$, and where $E[u(.):]$ is of the form $f(w^*_R;w^R_R w)$, with $f$ increasing in its first argument and decreasing in its second argument.

For further reference it is useful to define the efficient set variables (see, e.g., Ingersoll (1987)):

$$A = \begin{bmatrix} 1 \end{bmatrix} R K^1; B = \begin{bmatrix} 1 \end{bmatrix} R K^1; \text{ and } C = \begin{bmatrix} 1 \end{bmatrix} R K^1.$$  

A mean-variance efficient portfolio $w^*$ is the solution to the problem

$$\max_{f,w} L = w^*_R w + \omega R R^1 w^* R K^1 (w^* R K^1);$$

where $\omega$ is the coefficient of risk aversion. From the first order conditions of this problem it follows that a portfolio $w^*$ is mean-variance efficient if there
exist scalars \( \varphi \) and \( \gamma \) such that

\[ w^\mu = \varphi \gamma \left( \begin{array}{c} 1 \\ \gamma \end{array} \right) = \gamma \left( \begin{array}{c} \frac{1}{R} \\ \frac{1}{R} \end{array} \right) \]  

(8)

Because of the restriction \( w^0 = 1 \), it also follows that \( \varphi = B \gamma A \gamma \), implying that each mean-variance efficient portfolio is uniquely determined when either \( \varphi \) or \( \gamma \) is known, unless \( \gamma = B = A \). It is straightforward to show that for a given mean-variance efficient portfolio \( w^\mu \), the Lagrange multiplier \( \gamma \) equals the expected return on the zero-beta portfolio of \( w^\mu \), i.e., the intercept of the line tangent to the mean-variance frontier at \( w^\mu \) (in mean-standard deviation space). Since \( B = A \), is the expected return on the global minimum variance (GMV) portfolio, this is the intercept of the asymptotes of the mean-variance frontier, but there are no lines tangent to the frontier originating at this point (see, e.g., Ingersoll (1987, p.86)).

To show the duality between mean-variance frontiers and volatility bounds, take \( \gamma (v) \) for a given \( v \), and choose a mean-variance efficient portfolio such that \( \gamma = 1 \gamma v \). It follows from (8) and (5) that

\[ w^\mu (v) = \frac{\frac{1}{R} \gamma}{B \gamma A} \left( \begin{array}{c} 1 \\ v \end{array} \right) = \frac{\frac{1}{R} \gamma}{A \gamma v B} \gamma (v) = \frac{1}{R} v \gamma (v); \]

which shows that the vector \( \gamma (v) \) is proportional to a mean-variance efficient portfolio with zero-beta return equal to \( 1 \gamma v \). Thus, each point on the volatility bound of stochastic discount factors, i.e., \( (v, \text{Var}[m(v)_{t+1}]) \) corresponds to a unique point on the mean-variance frontier, \( \{1 \gamma v\} \), and each coefficient vector \( \gamma (v) \) corresponds to a unique \( w^\mu (v) \). The only exception to this result is the case where \( \gamma (v) = 0 \), which is the case if \( v = A = \gamma B \), or equivalently, \( \gamma = B = A \). As already noted, this is the case where the zero-beta return equals the expected return on the global minimum variance portfolio (see also Hansen and Jagannathan (1991)). The duality between the mean-variance frontier of \( R_{t+1} \) and the volatility bound derived from \( R_{t+1} \) can also be seen directly from (5) and (8). Comparing the coefficients \( \gamma (v) \) for the minimum variance stochastic discount factor in (5) and the portfolio weights \( w^\mu \) in (8) for \( \gamma = 1 \gamma v \), it can be seen that the coefficients \( \gamma (v) \) are proportional to the portfolio weights \( w^\mu \), where the coefficient of proportionality is equal

\[ 2 \text{More precisely, these are the minimum variance portfolios, i.e., the portfolios that have minimum variance for a given expected return. The mean-variance efficient portfolios, i.e., the portfolios that also have maximum expected return for a given variance, require in addition that } \varphi, 0. \]
to $j \neq \varphi$, i.e., $w^\varphi = (j \neq \varphi)'(v)$. In Appendix A we show graphically which points on the volatility bound correspond to points on the mean-variance frontier.

Summarizing, finding stochastic discount factors that have the lowest variance of all stochastic discount factors that price a set of asset returns $R_{t+1}$ correctly is tantamount to finding mean-variance efficient portfolios for these same assets $R_{t+1}$. In the remainder of this paper we will study the effects of adding new assets to the set of assets available to investors. Although most of the results will be stated in terms of mean-variance frontiers and mean-variance efficient portfolios, it should be kept in mind that there is always a dual interpretation in terms of volatility bounds.

3 Mean-variance spanning and intersection

In the previous section we considered the volatility bounds and mean-variance frontiers that can be derived from a given set of $K$ assets with return vector $R_{t+1}$. Suppose now that an investor takes an additional set of $N$ assets with return vector $r_{t+1}$ into account in his portfolio problem. The question we are interested in is under what conditions mean-variance efficient portfolios derived from the set of returns $R_{t+1}$ are also mean-variance efficient for the larger set of $K + N$ assets $(R_{t+1}; r_{t+1})$. This problem was addressed in the seminal paper of Huberman and Kandel (1987). If there is only one value of $\varphi$ or $\varphi'$ for which mean-variance investors can not improve their mean-variance efficient portfolio by including $r_{t+1}$ in their investment set, the mean-variance frontiers of $R_{t+1}$ and $(R_{t+1}; r_{t+1})$ have exactly one point in common, which is referred to as intersection. In this case we will say that the mean-variance frontier of $R_{t+1}$ intersects the mean-variance frontier of $(R_{t+1}; r_{t+1})$, or simply that $R_{t+1}$ intersects $(R_{t+1}; r_{t+1})$. If there is no mean-variance investor that can improve his mean-variance efficient portfolio by including $r_{t+1}$ in his investment set, the mean-variance frontiers of $R_{t+1}$ and $(R_{t+1}; r_{t+1})$ coincide, which is referred to as spanning. In this case we will say that (the mean-variance frontier of) $R_{t+1}$ spans (the mean-variance frontier of) $(R_{t+1}; r_{t+1})$.

As suggested by the previous section, and as shown by Ferson, Foerster, and Keim (1993), DeSantis (1994), Ferson (1995) and Bekaert and Urias (1996), the concept of mean-variance spanning and intersection has a dual interpretation in terms of volatility bounds. In terms of volatility bounds mean-variance spanning means that the volatility bound derived from the
returns \( R_{t+1} \) is the same as the bound derived from \( (R_{t+1}; r_{t+1}) \). Therefore, the minimum variance stochastic discount factors for \( R_{t+1} \), \( m_r(v)_{t+1} \), are also the minimum variance stochastic discount factors for \( (R_{t+1}; r_{t+1}) \), and the asset returns \( r_{t+1} \) do not provide information about the necessary volatility of stochastic discount factors that is not already present in \( R_{t+1} \). As will be shown formally below, mean-variance intersection is equivalent to saying that the volatility bounds derived from \( R_{t+1} \) and \( (R_{t+1}; r_{t+1}) \) have exactly one point in common. Thus, in case of intersection there is exactly one value of \( v \) for which the minimum variance stochastic discount factor does not change, whereas for all other values of \( v \) it does.

In infinite samples it will in general be the case that adding assets causes a shift in the estimated mean-variance frontier and the estimated volatility bound. This shift may very well be the result of estimation error however, and the main question is whether the observed shift is too large to be attributed to chance. Therefore, to answer the question whether or not the observed shift in the mean-variance frontier is significant in statistical terms, in this section we will also show how regression analysis can be used to test for spanning and intersection.

### 3.1 Spanning and intersection in terms of mean-variance frontiers

To state the problem formally, the hypothesis of mean-variance intersection means that there is a portfolio \( w^\alpha \) which is mean-variance efficient for the smaller set \( R_{t+1} \) and which is also mean-variance efficient for the larger set \( (R_{t+1}; r_{t+1}) \). In the sequel, variables that refer to the smaller set \( R_{t+1} \) \((r_{t+1})\) will be referred to with a subscript \( R \) \((r)\), or with their dimension \( K \) \((N)\), whereas variables that refer to the larger set \( (R_{t+1}; r_{t+1}) \), will not have any subscript or will have their dimension as subscript, \( K + N \). Thus, \( w_R \) is a \( K \)-dimensional vector with portfolio weights for the assets in \( R_{t+1} \), and \( w \) is a \((K + N)\)-dimensional vector with portfolio weights for all the available assets \( (R_{t+1}; r_{t+1}) \). The hypothesis of mean-variance intersection comes down to the statement that there exists a mean-variance efficient portfolio \( w^\alpha \) of the form

\[
w^\alpha = \begin{bmatrix} w_R \\ 0_N \end{bmatrix};
\]
i.e., there exist scalars \( \circ \) and \( \cdot \), such that

\[
\begin{equation}
1 \ i \ \cdot \ k + N = \circ \ \tilde{A} \ \bar{w}^R \end{equation}
\]

If such a portfolio \( w^R \) exists, there is one point on the mean-variance frontier of \( R_{t+1} \) that also lies on the mean-variance frontier of \((R_{t+1}; r_{t+1})\). Using obvious notation, \( \circ \) consists of two subvectors \( \circ \ R \) and \( \circ \ r \), and \( \circ \ \tilde{A} \) consists of submatrices \( \circ \ RR, \circ \ Rr, \circ \ rR, \) and \( \circ \ rr \). The first \( K \) rows of (11) imply that

\[
\begin{equation}
\begin{align*}
1 \ R \ i \ \cdot \ k \ &= \ \circ \ \circ \ RR \ \bar{w}^R \\
\bar{w}^R &= \circ \ \circ \ \circ \ R \ \bar{R} (1 \ R \ i \ \cdot \ k)
\end{align*}
\end{equation}
\]

Equation (12) simply says that \( w^R \) is indeed mean-variance efficient for the smaller set \( R_{t+1} \).

The next step is to derive the restrictions on the distribution of \( R_{t+1} \) and \( r_{t+1} \) that are equivalent to mean-variance intersection. In order to do so, substitute (12) in the last \( N \) rows of (11) to obtain:

\[
\begin{equation}
\begin{align*}
1 \ r \ i \ \cdot \ k \ &= \ \circ \ \circ \ Rr \ \bar{w}^R \\
\bar{w}^R &= \circ \ \circ \ \circ \ \circ \ RR (1 \ R \ i \ \cdot \ k)
\end{align*}
\end{equation}
\]

with \( \circ \ \circ \ \circ \ RR \). Thus, if there is a portfolio that is mean-variance efficient for the smaller set \( R_{t+1} \) that is also mean-variance efficient for the larger set \((R_{t+1}; r_{t+1})\), there must exist an \( \cdot \) such that the restriction in (13) holds. It follows immediately from the derivation above that this \( \cdot \) is the zero-beta return that corresponds to the portfolio \( w^R \) (and \( w^\circ \)).

If there is mean-variance spanning then all mean-variance efficient portfolios \( w^\circ \) must be of the form (10), i.e., (11) must be true for all values of \( \cdot \) and the corresponding \( \circ \)'s. Going through the same steps, if (11) must hold for any \( \cdot \), (13) must hold for any \( \circ \), and this can only be the case if

\[
\begin{equation}
\begin{align*}
1 \ r \ i \ \cdot^{-1} R &= 0 \quad \text{and} \quad \circ \ i \ \cdot \ N = 0
\end{align*}
\end{equation}
\]

which are the restrictions imposed by the hypothesis of spanning. If these restrictions on the distribution of \( R_{t+1} \) and \( r_{t+1} \) hold, every point on the mean-variance frontier of \( R_{t+1} \) is also on the mean-variance frontier of \((R_{t+1}; r_{t+1})\) and the two frontiers coincide.
3.2 Spanning and intersection in terms of volatility bounds

In the previous section we defined mean-variance spanning and intersection from the properties of mean-variance efficient portfolios and we derived the equivalent restrictions on the distribution of asset returns, which have previously been derived by Huberman and Kandel (1987). In this section we analyze mean-variance intersection and spanning from the properties of minimum variance stochastic discount factors that price the assets in \( R_{t+1} \) and \( (R_{t+1}; r_{t+1}) \) correctly and we show that this imposes the same restrictions on the distribution of the asset returns. In terms of volatility bounds, the hypothesis of intersection is that there is a value of \( v \) such that the minimum variance stochastic discount factor for \( R_{t+1} \), i.e., \( m_R(v)_{t+1} \), is also the minimum variance stochastic discount factor for the larger set \( (R_{t+1}; r_{t+1}) \). The discount factor \( m_R(v)_{t+1} \) as defined by (4) and (5) is the minimum variance stochastic discount factor for this larger set if it also prices \( r_{t+1} \) correctly. If \( m_R(v)_{t+1} \) prices both \( R_{t+1} \) and \( r_{t+1} \) correctly, the difference between \( m_R(v)_{t+1} \) and any other \( M(v)_{t+1} \) that prices \( R_{t+1} \) and \( r_{t+1} \) correctly is orthogonal to \( R_{t+1} \) and \( r_{t+1} \), implying that \( m_R(v)_{t+1} \) must have the lowest variance among all stochastic discount factors \( M(v)_{t+1} \), by the same reasoning that leads to (6).

Thus, the hypothesis of intersection for volatility bounds can be stated as:

\[
9v \text{ s.t. } E[r_{t+1}m_R(v)_{t+1}] = \mu_N; \quad (15)
\]

To show that this hypothesis imposes the same restrictions on the distribution of \( R_{t+1} \) and \( r_{t+1} \) as in (13), substitute (4) and (5) into (15):

\[
E[r_{t+1}(v + (R_{t+1} i 1_R)^\top \hat{\Sigma}_R i 1_R (\hat{\mu} i v^\top R))] = \mu_N; \quad (16)
\]

Dividing both sides of (16) by \( v \) shows that the hypothesis of intersection in terms of volatility bounds indeed implies the same restrictions as the hypothesis of intersection in terms of mean-variance frontiers, if we choose \( \lambda = 1 = \nu \). This could be expected beforehand, since from the duality between mean-variance frontiers and volatility bounds in (9) we already knew that the vector \( R(v) \) that defines \( m_R(v)_{t+1} \), is proportional to a mean-variance efficient portfolio with zero-beta return \( \lambda = 1 = \nu \). The hypothesis that \( w^a \) is
of the form \((w_R^0 0_R^0)^0\) is therefore equivalent the hypothesis that \(\gamma(v)\) is of the form \((\gamma(v)^0 0_R^0)^0\).

By the same logic, the hypothesis of spanning in terms of volatility bounds, requires that \(m_R(v) \) prices the returns \(r_{t+1}\) for all values of \(v\):

\[
E[r_{t+1}m_R(v)] = \mu; \quad \forall v; \quad (17)
\]

since in that case the entire volatility bound derived from \((R_{t+1}^0 r_{t+1}^0)\) coincides with the volatility bound derived from \((R_{t+1}^0)\) only. This requirement implies that (16) holds for all values of \(v\), and this can only be the case if the restrictions in (14) hold.

### 3.3 Intersection and mean-variance efficiency of a given portfolio

A question that is of obvious interest both from a portfolio choice perspective and from an asset pricing perspective, is the question whether or not a given portfolio \(w^p\) is mean-variance efficient or not. From a portfolio choice perspective, an investor will be interested in whether or not his portfolio has the desired properties of a mean-variance efficient portfolio. From an asset pricing perspective, the frequently analyzed question is, e.g., whether or not the market portfolio is mean-variance efficient as the CAPM predicts. Alternative asset pricing models may identify other portfolios as being mean-variance efficient. For instance, in the Consumption-CAPM the portfolio that mimics aggregate per-capita consumption is mean-variance efficient and the Intertemporal-CAPM implies that a combination of the market portfolio and the portfolios hedging changes in the investment-opportunity set is mean-variance efficient.

Denote the return on some portfolio \(w^p\) by \(R_{t+1}^p\) and its expectation by \(\mu^p\). The question whether or not \(w^p\) is mean-variance efficient with respect to the \(N + 1\) assets \((R_{t+1}^p; r_{t+1}^p)\), is obviously a special case of the question whether or not there is mean-variance intersection with \(K = 1\) and \(R_{t+1} = R_{t+1}^p\), since intersection in this case simply means that the portfolio \(w^p\) is on the mean-variance frontier of \((R_{t+1}^p; r_{t+1}^p)\). Therefore, if \(w^p\) is mean-variance efficient for the set \((R_{t+1}^p; r_{t+1}^p)\), the following restrictions on the distribution of \(R_{t+1}^p\) and \(r_{t+1}\) should hold:

\[
\gamma_r = \mu^p + \rho^p (1^p \epsilon^p); \quad (18)
\]
where $\mathbf{\bar{P}}$ is the $N$-dimensional vector $\text{Cov}[r_{t+1};R_{t+1}^p] = \text{Var}[R_{t+1}^p]$, and $1^p = E[R_{t+1}^p]$. When testing for mean-variance efficiency, $R_{t+1}^p$ is usually the return on a portfolio of $R_{t+1}$.

What we want to establish in this section however, is that the hypothesis that the mean-variance frontier of $R_{t+1}$ ($K > 1$) intersects the frontier of $(R_{t+1}^p; r_{t+1})$ at a given value of $\gamma = 1 = \nu$, is tantamount to the hypothesis that the portfolio $w^p_R$ that is mean-variance efficient for $R_{t+1}$ and that has $\gamma$ as its zero-beta rate is also mean-variance efficient with respect to $(R_{t+1}; r_{t+1})$. Denote the return on $w^p_R$ as $R_{t+1}^p$ and its expectation as $\mu^p_R$. Recall that the portfolio $w^p_R$ is given by the first $K$ rows of (11)

$$w^p_R = \mathbf{\sigma} \gamma \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1} \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1} \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1}$$

from which

$$w^p_R = \mathbf{\sigma} \gamma \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1}$$

Substituting these relations into (11) and defining $\mathbf{\bar{R}}^p = \text{Cov}[r_{t+1}; R_{t+1}^p] = \text{Var}[R_{t+1}^p]$,

results in

$$0 = \left(\gamma \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1} \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1}\right) - \left(\gamma \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1}\right)^\gamma$$

These are the same restrictions as (18) for $w^p = w^p_R$. Thus, the hypothesis of intersection indeed implies the same restrictions on the distribution of $R_{t+1}$ and $r_{t+1}$ as the hypothesis that $w^p_R$ is mean-variance efficient with respect to $R_{t+1}$.

### 3.4 Testing for spanning and intersection

So far we derived the restrictions implied by the hypotheses of mean-variance intersection and spanning for the distribution of $R_{t+1}$ and $r_{t+1}$. Huberman and Kandel (1987) showed how regression analysis can be used to test these hypotheses. To see how regression analysis can be used to test for intersection, start from (13):

$$\gamma \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1} \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1} \mathbf{\bar{R}} \mathbf{\bar{R}}^{-1}$$

Replacing the expected returns $\gamma$ and $\gamma$ with realized returns $r_{t+1}$ and $R_{t+1}$, gives the regression

$$r_{t+1} = \mathbf{\sigma} + \mathbf{\bar{R}} r_{t+1} + \mathbf{\bar{R}} r_{t+1},$$

(20)
with \( \beta = 1_r \), \( \bar{R}_t = u_{r:t+1}^{-1} R_{t+1} \), and \( u_{r:t+1}^{-1} R_{t+1} \). It can readily be checked that under the null hypotheses of spanning and intersection \( \text{Cov}[^t; R_{t+1}] = 0 \). Notice that \( \beta \) is an \( N \)-dimensional vector of intercepts, \( \bar{\beta} \) is an \( N \times K \)-dimensional matrix of slope coefficients, and \( ^t; \) is an \( N \)-dimensional vector of error terms. The restrictions imposed by the hypothesis of intersection in (13) can now be stated as

\[
\beta_i (\bar{\beta}_i \bar{\Pi}_k) = 0; \quad (21)
\]

With intersection there are two cases of interest. First, we may be interested in testing for intersection for a given value of the zero-beta rate \( \beta \). In that case the restrictions in (21) should hold for this specific value of \( \beta \), which is a set of linear restrictions. In the sequel we will mainly be interested in this case. Second, the interest may be in the question whether there is intersection at some unknown point of the frontier, i.e., for some unknown value of \( \beta \). In that case the hypothesis is that there exists some \( \beta \) such that the restrictions in (21) hold. This hypothesis can be stated as

\[
\beta_i (1_{i \bar{\beta}_i} \bar{\Pi}_k) = \beta_j (1_{j \bar{\beta}_j} \bar{\Pi}_k); \quad i, j = 1, \ldots, N;
\]

where \( 1_i \) is the \( i \)-th row of \( \bar{\beta} \). Thus, the hypothesis that there is intersection at some point of the frontier imposes a set of nonlinear restrictions on the regression parameters in (20). Notice that given estimates of \( \beta \) and \( \bar{\beta}_i \), an estimate of the zero-beta rate for which there is intersection can be obtained from \( \beta_i (1_{i \bar{\beta}_i} \bar{\Pi}_k) \). Also note, that testing whether there is intersection at some unknown point of the frontier only makes sense if \( N \geq 2 \), since there is always intersection if \( N = 1 \). (Because there is always one efficient portfolio for which the weight in the new asset is zero.)

Recall that the hypothesis of spanning implies that (21) holds for all values of \( \beta \). Therefore, going through the same steps, the restrictions imposed by the hypothesis of spanning can be stated as

\[
\beta = 0 \quad \text{and} \quad \bar{\Pi}_k i \bar{\Pi}_N = 0. \quad (22)
\]

The restrictions in terms of the regression model in (20) are intuitively very clear. For instance, the spanning restrictions in (22) state that if there is spanning, then each return of the additional assets, \( r_{i:t+1}, i = 1, 2, \ldots, N \), can be written as the return of a portfolio of the benchmark assets \( \bar{\beta}_i R_{t+1}, \bar{\Pi}_k = 1 \), plus an error term \( ^t; i \bar{\Pi}_N \) which has expectation zero and which
is orthogonal to the returns $R_{t+1}$. Since such an asset can only add to the variance of portfolios of $R_{t+1}$, and not to the expected return, mean-variance optimizing agents will not include such an asset in their portfolio. A similar interpretation holds for the intersection restrictions.

If the returns series $R_{t+1}$ and $r_{t+1}$ are stationary and ergodic, consistent estimates of the parameters $\beta$ and $\gamma$ in (20) are easily obtained using OLS. In writing down the test statistics for (21) and (22), it is convenient to use a different specification of (20), in which all the coefficients $\beta$ and $\gamma$ are stacked into one big vector:

$$r_{t+1} = I_N - 1 R_{t+1}^0 b + "t+1;$$  \hspace{1cm} (23)

where $b = \text{vec} \begin{pmatrix} \beta \\ 0 \end{pmatrix}$, a $(K + 1)N$-dimensional vector. If $\hat{b}$ is the OLS estimate of $b$ and $\hat{Q}$ is a consistent estimate of the asymptotic covariance matrix of $\hat{b}$, the hypotheses of intersection and spanning can be tested using a standard Wald test. Defining

$$H(\hat{\cdot})_\text{int} = I_N - 1 \begin{pmatrix} \beta \\ 0 \end{pmatrix} \text{ and } h(\hat{\cdot})_\text{int} = H(\hat{\cdot})_\text{int} \hat{b}_i \hat{Q} \hat{b}_i,$$  \hspace{1cm} (24a)

the Wald test-statistic for intersection can be written as

$$\lambda^\text{int}_W = h(\hat{\cdot})_\text{int}^0 H(\hat{\cdot})_\text{int} \hat{Q} H(\hat{\cdot})_\text{int}^0 \hat{b}_i \hat{Q} \hat{b}_i.$$  \hspace{1cm} (25)

Similarly, defining

$$H(\hat{\cdot})_\text{span} = I_N - 1 \begin{pmatrix} \beta \\ 0 \end{pmatrix} \text{ and } h(\hat{\cdot})_\text{span} = H(\hat{\cdot})_\text{span} \hat{b}_i \hat{Q} \hat{b}_i,$$  \hspace{1cm} (26a)

the Wald test-statistic for spanning can be written as

$$\lambda^\text{span}_W = h(\hat{\cdot})_\text{span}^0 H(\hat{\cdot})_\text{span} \hat{Q} H(\hat{\cdot})_\text{span}^0 \hat{b}_i \hat{Q} \hat{b}_i.$$  \hspace{1cm} (27)

Under the null hypotheses and standard regularity conditions, the limit distribution of $\lambda^\text{int}_W$ will be $\tilde{A}^2_N$ and the limit distribution of $\lambda^\text{span}_W$ will be $\tilde{A}^2_{2N}$. The test statistics in (25) and (27) have interesting economic interpretations.
in terms of performance measures. The relationship between tests for intersection and spanning and performance evaluation will be discussed in detail in Section 5.3.

Chen and Knez (1996) and Hall and Knez (1995) propose a test for intersection that is based on (15). Define the deviation from the equality in (15) to be \( \hat{\varepsilon}(v) \):

\[
\hat{\varepsilon}(v) = E[m_R(v_t r_t) | V]:
\]

(28)

In Section 5.1 we will interpret \( \hat{\varepsilon}(v) \) scaled by \( v \) as a generalization of the well-known Jensen measure. Given an estimate of the parameters \( 'R(v) \) using the sample equivalent of (5):

\[
b_R(v) = \frac{1}{T} \sum_{t=1}^{T} (R_t - \bar{R})(R_t - \bar{R})' \bar{R}^{-1} \bar{V}:
\]

with \( \bar{R} \) the sample mean of \( R_t \), define \( \hat{b}(v)_t \) as

\[
\hat{b}(v)_t = r_t(v + b_R(v)'(R_t - \bar{R}))' \bar{V}:
\]

A test for the hypothesis of intersection, \( \hat{\varepsilon}(v) = 0 \), can now be based on

\[
\hat{\chi}^{\text{int}}_{\text{CK}} = \frac{1}{T} \sum_{t=1}^{T} b(v)_t' \varphi \text{ar}[b(v)_t] \frac{1}{T} \sum_{t=1}^{T} b(v)_t;
\]

where the estimate \( \varphi \text{ar}[b(v)_t] \) can for instance be obtained using the method suggested by Newey and West (1987). The limit distribution of the test-statistic \( \hat{\chi}^{\text{int}}_{\text{CK}} \) is also \( \chi^2_N \). Since for \( \beta = 1 = v \), we have

\[
\frac{1}{T} \sum_{t=1}^{T} b(v)_t = \frac{1}{T} \sum_{t=1}^{T} (R_t - \bar{R})' R_t + \bar{R}^{-1} \bar{V} \hat{b}(v)_t' \bar{V} - \frac{1}{T} \sum_{t=1}^{T} b(v)_t'
\]

\[
= \hat{b} + \hat{b}_k' \bar{V}:
\]

it follows that

\[
\frac{1}{T} \sum_{t=1}^{T} b(v)_t = \varphi (\beta)_\text{int} \hat{b}_i + h(\beta)_\text{int}:
\]

and that the only difference in the Wald test-statistic in (25) and the statistic proposed in (29) is the way in which the covariance matrix is estimated.

A disadvantage of the test originally proposed by Chen and Knez (1996) is that they test for intersection for a very specific stochastic discount factor.
which corresponds to the minimum second moment portfolio. This discount factor can be found by projecting the kernel \( M_{t+1} \) on the asset returns only, excluding the constant. The corresponding portfolio on the mean-variance frontier is the one with the minimum second moment among all portfolios on the frontier, and can graphically be found as the tangency point between the mean-variance frontier and a circle with its centre at the origin. The problem with this portfolio is that it is located at the inefficient part of the frontier, implying that the test used by Chen and Knez (1995) is for intersection at an efficient portfolio. Therefore it is economically not very interesting, unless a risk free asset is included. Since in the test statistic in (29) the discount factor \( m_R(v)_{t+1} \) results from a projection of \( M_{t+1} \) on \( R_{t+1} \) plus a constant, this test allows us to test for intersection at any mean-variance efficient portfolio, so this test does not suffer from the problem of the test originally suggested by Chen and Knez. Dahlquist and Söderlind (1999), who use the test proposed by Chen and Knez to evaluate the performance of Swedish mutual funds, also acknowledge this problem and add a constant to the set \( R_{t+1} \) such that the conditional mean of \( m_r(v)_{t+1} \), equals one over the risk free rate, i.e., \( v_t = 1/R_{f,t} \).

The distinction between the Wald tests in (25) and (27) on the one hand and the tests proposed by Chen and Knez in (29) is similar to the distinction between tests based on the (traditional) regression methodology and on the SDF methodology as discussed in Kan and Zhou (1999). Their simulations suggest that in small samples tests based on the regression methodology have better size and power properties than tests based on the SDF methodology, which indicates that the test in (25) may be preferred to (29).

Alternative tests for the hypotheses of intersection and spanning are suggested, e.g., by Huberman and Kandel (1987), who propose a likelihood ratio test, and by Snow (1991) and DeSantis (1995), who propose a Generalized Method of Moments (GMM) procedure. This latter procedure is also identical to the region subset test suggested by Hansen, Heaton and Luttmer (1995) which is equivalent to a test for intersection. A comparison of the small sample properties of various test procedures can be found in Bekaert and Urias (1996). Their small sample results suggest that the likelihood test for spanning as proposed by Huberman and Kandel has better power properties than the GMM-based tests, while it also has a size distortion that is in most cases not worse than for the GMM-based tests. The GMM-based test or region subset test is based on the observation that under the null hypotheses of spanning or intersection, the kernel that prices \( R_{t+1} \) and \( f_{t+1} \)
correctly is of the form
\[ m(v)_{t+1} = v + 'R(v)q_{R_{t+1}} + 'r(v)q_{r_{t+1}}. \]

Given that \( 'r(v) = 0 \), a GMM-estimate of the \( K \) parameters in \( 'R(v) \) can be obtained by using the \( K + N \) sample moments
\[ g_T('R(v)) = \frac{1}{T} \sum_{t=1}^{T} \left( \hat{A}_{R_t}^{-1} \left( \frac{1}{T} \sum_{t=1}^{T} \left( v + 'R(v)q_{R_t} + 'r(v)q_{r_t} \right) \right) \right) \]

A consistent estimate of \( 'R(v) \) can therefore be obtained by solving
\[ \min_{'R(v)} g_T('R(v))W_Tg_T('R(v)) = J_T('R(v)); \]

where \( W_T \) is a symmetric nonsingular weighting matrix. Notice that the GMM-estimate of the \( K \) parameters in \( 'R(v) \) obtained from (30) is based on \( K + N \) moment restrictions. The \( N \) overidentifying restrictions are derived from the hypothesis that \( m_R(v)_{t+1} \) must also price the \( N \) additional assets \( r_{t+1} \). Intersection for a given value of \( v \) can now be tested by using the fact that under the null-hypothesis and regularity conditions \( T_j(a_R(v)) \) is asymptotically \( \chi^2_N \)-distributed. Since spanning implies that (15) holds for (at least) two different values of \( v \), the GMM-based test can easily be extended by estimating two vectors \( 'R(v_1) \) and \( 'R(v_2) \) simultaneously \( (v_1 \neq v_2) \) using (30). In this case there are \( 2K \) parameters to be estimated with \( 2(K + N) \) moment conditions. The test for spanning is therefore a test for the \( 2N \) overidentifying restrictions and will asymptotically be \( \chi^2_{2N} \)-distributed under the null-hypothesis of spanning.

4 Testing for spanning and intersection with conditioning information

The purpose of this section is to incorporate conditioning information in tests for intersection and spanning. Until now we assumed that returns are independently and identically distributed (i.i.d.). However, there is ample evidence that asset returns are to some extent predictable. For instance, stock and bond returns can be predicted from variables like lagged returns,
dividend yields, short term interest rates, and default premiums (see, e.g., Ferson (1995)) and futures returns can be predicted from hedging pressure variables (see e.g. DeRoon, Nijman and Veld (2000)) as well as from the spread between spot and forward prices (see, e.g., Fama (1984)). Kirby (1998) analyzes whether predictability of security returns is consistent with rational asset pricing. He shows that the covariance between the pricing kernel implied by an asset pricing model and conditioning variables, restricts the slope coefficients in a regression of security returns on those same conditioning variables. In Section 4.1 we will show how conditional information can be used in a straightforward way by using scaled returns (see, e.g., Cochrane (1997) and Bekaert and Urias (1996)). Although this is a fairly general and intuitive way of incorporating conditional information, a disadvantage of this method is that the dimension of the estimation and testing problem increases quickly. In Section 4.2 we show that this problem can be circumvented if it is assumed that variances and covariances are constant, while expected returns are allowed to vary over time, although this assumption is not in accordance with most equilibrium models and with the empirical evidence regarding time-varying second moments. Using this simplifying assumption however, it is shown that the conditioning variables can easily be accounted for by using them as additional regressors. The restrictions for the intersection and spanning hypotheses then become similar to the restrictions in the i.i.d. case. This way of incorporating conditional variables also has the additional advantage that the regression estimates indicate under what economic circumstances, i.e., for what values of the conditioning variables, intersection and spanning can or cannot be rejected. Finally, in Section 4.3 we will discuss the use of conditioning variables as, e.g., in Shanken (1990) and Ferson and Schadt (1996).

4.1 Incorporating conditional information using scaled returns

Suppose that \( \mathbf{Z}_t \) is an \((L+1)\)-dimensional vector of instruments that has predictive power for \( R_{t+1} \) and \( r_{t+1} \), and define the \( L \)-dimensional vector \( \mathbf{Z}_t \) as \( \mathbf{Z}_t \left( \mathbf{z}_0^{(1)} + \mathbf{z}_0 \right) \). A common way to use these instruments is to look at scaled returns: \( \mathbf{Z}_t - R_{t+1} \). If \( M_{t+1} \) is a valid stochastic discount factor, then from (1) we have:

\[
E[M_{t+1}(\mathbf{Z}_t - R_{t+1}) | \mathbf{I}_t] = \mathbf{Z}_t - \mathbf{q}:
\]
Taking unconditional expectations, this yields

$$E [M_{t+1}(Z_t - R_{t+1})] = E [Z_t - \mu_k];$$

(31)

Thus, the scaled return $Z_i; R_{j; t+1}$ has an average price equal to $E [Z_i; t]$. The
scaled returns can be interpreted as the payoffs of a strategy where each
period an amount equal to $Z_i$ dollars is invested in a security, yielding
a payoff equal to $Z_i; R_{j; t+1}$. Therefore, we can also think of $Z_t - R_{t+1}$ as
the returns on managed portfolios (see, e.g., Cochrane (1997)). By allowing
for such managed portfolios, we take into account that investors may use
dynamic strategies, based on the realized values of $Z_t$. In effect this increases
the set of available assets by a factor $L$ (i.e., from $K$ to $K \cdot L$).

To simplify notation, denote the $(L \times K)$-dimensional vector $Z_t - R_{t+1}$
by $R_{t+1}$. Also, denote the $(L \times K)$-dimensional vector $E [Z_t - \mu_k]$ by $q_k$. For
further reference, $r_{t+1}$ and $q_k$ are defined in a completely analogous way and
we use a superscript $Z$ for all variables and parameters that correspond to
$R_{t+1}$ and $r_{t+1}$. Valid stochastic discount factors $M_{t+1}$ now have to satisfy

$$E [M_{t+1}^Z R_{t+1}^Z] = q_k;$$

(32)

As shown by Bekaert and Urias (1996), following the same line of reasoning
as in Sections 2.1 and 2.2, it is straightforward to show that the minimum
variance stochastic discount factor with expectation $v$ is given by

$$m_{R}^{Z}(v)_{t+1} = v + Z(v)^{T}R_{t+1}^{Z}1^{Z}_{R};$$

$$Z(v) = (R_{Z}^{Z}R_{R})^{-1}1^{Z}(q_k) + \lambda^{Z}Z_{R};$$

(33)

This expression for the volatility bound is a straightforward generalization of
the one given in (4) and (5). The restrictions imposed by the hypotheses of
intersection and spanning also turn out to be very similar to the ones given
in previous sections, as we will see below.

Thus, conditioning information can be incorporated by including managed
portfolios, the returns of which depend on the conditioning variables. If
there is to be conditional intersection or spanning of $r_{t+1}$ by $R_{t+1}$, the uncondi-
tional volatility bound (or mean-variance frontier) of $R_{t+1}$ must intersect
or span the volatility bound (or mean-variance frontier) of $(R_{t+1}; r_{t+1})$. The
interest is therefore in the returns $R_{t+1}$ and $r_{t+1}$ themselves plus the returns
on all the managed portfolios. Intersection or spanning is equivalent to

$$E [r_{t+1}^{Z}m_{R}^{Z}(v)_{t+1}] = q_k;$$

(34)
for one value of \( v \) or for all values of \( v \) respectively. To see which restrictions these hypotheses imply, substitute (33) into (34) to obtain

\[
\begin{pmatrix} Z_r & -Z_i \end{pmatrix} v + \begin{pmatrix} -Z q_K & q_N \end{pmatrix} = 0; \tag{35}
\]

for intersection, and

\[
\begin{pmatrix} Z_r & -Z_i \end{pmatrix} = 0; \quad \text{and} \quad \begin{pmatrix} -Z q_K & q_N \end{pmatrix} = 0; \tag{36}
\]

for spanning. Here \( Z = \begin{pmatrix} Z_r \end{pmatrix} \) is a \((L \times N) \times (L \times K)\) matrix with slope coefficients from a regression of \( r_{t+1}^Z \) on \( R_{t+1}^Z \) plus a constant. These restrictions are also given in Bekaert and Urias (1996). Regressing \( r_t^Z \) on \( R_t^Z \) to incorporate conditioning information is very similar to the approach to be discussed in Section 4.3, where the regression parameters \( \beta \) and \( \gamma \) are time varying. In that section we will assume that the mean returns and the (co)variances are functions of the instruments that can be linearized using a Taylor series approximation, leading to a similar regression as in the case discussed here. Therefore, the use of scaled returns can also be motivated as a convenient way of dealing with time-varying means and variances.

The similarity with the case in which there was no conditioning information is obvious. The only difference in the restrictions is that in (35) and (36) we have \( \begin{pmatrix} -Z q_K & q_N \end{pmatrix} \) instead of \( \begin{pmatrix} -q_K & q_N \end{pmatrix} \). The fact that \( q_K \) and \( q_N \) enter the restrictions reflects the fact that \( R_{t+1}^Z \) and \( r_{t+1}^Z \) are not really returns, in the sense that their current prices are not necessarily equal to one. The average prices of \( R_{t+1}^Z \) and \( r_{t+1}^Z \) are instead given by \( q_K \) and \( q_N \). The average cost of the managed portfolios with payoff vector \( r_{t+1}^Z \) is given by the vector \( q_N \), and the cost of the mimicking portfolios from \( R_{t+1}^Z \) is given by \( -Z q_K \). The interpretation of the restrictions given in Section 3.4 is therefore still valid.

The main disadvantage of this way of incorporating conditioning information is that the number of parameters to be estimated as well as the number of restrictions to be tested grows rapidly with the number of instruments \( L \). The number of exogenous variables equals \( K \times L \) and the number of restrictions to be tested equals \( N \times L \) for the hypothesis of intersection, and \( 2N \times L \) for the hypothesis of spanning. This is the case because for each new instrument there are \( K \) new managed portfolios to be considered for the assets in \( R_{t+1} \) and \( N \) additional managed portfolios for the assets in \( r_{t+1} \).

This problem can at least partially be circumvented if we are willing to assume a more specific form of predictability. Specifically, in the next section
we make the assumption that only the expected returns of $R_{t+1}$ and $r_{t+1}$ depend linearly on the instruments $z_t$, whereas all variances and covariances are constants. In Section 4.3 the slope coefficients are assumed to depend linearly on the instruments, which also allows for a straightforward way of incorporating conditional information in the regression framework to test for intersection and spanning.

### 4.2 Expected returns linear in the conditional variables

In this section we assume that there is a specific form of predictability, which allows us to incorporate conditioning information in a straightforward way in the regression framework for spanning and intersection. The assumption made is that expected returns are linear in the conditional variables and that returns are conditionally homoskedastic. This way of incorporating conditioning information is used in Harvey (1989), as well as, for instance, in Campbell and Viceira (1998) and DeRoon, Nijman and Werker (1998). The assumption we make is that

$$E_t[R_{t+1}] = c_0^R Z_t;$$
$$E_t[r_{t+1}] = c_0^r Z_t;$$

and the variances and covariances of $R_{t+1}$ and $r_{t+1}$ conditional on $Z_t$ are given by

$$\text{Var}[R_{t+1} \mid Z_t] = \sigma_R^2, \quad \text{Var}[r_{t+1} \mid Z_t] = \sigma_r^2, \quad \text{Cov}[r_{t+1}; R_{t+1} \mid Z_t] = -\sigma_{Rr}.$$

Starting from (1), the minimum variance stochastic discount factor, conditional on $Z_t$, is given by

$$m_R(\nu_t)_{t+1} = \nu_t + (\nu_t)^{\prime} (R_{t+1} \mid E_t[R_{t+1}]);$$
$$m_N(\nu_t)_{t+1} = \nu_t \nu_t^{\prime} (\nu_t)_{t+1} E_t[R_{t+1}];$$

Notice that since the projection of the kernel on the asset returns is now conditional on $Z_t$, we explicitly allow for time variation in the coefficients $(\nu_t)_t$, as well as in $\nu_t$, the conditional expectation of the stochastic discount factor. Also note that in describing the conditional mean-variance frontier or volatility bound we still can use $\nu_t$ as a free parameter.

If there is intersection, $m_R(\nu_t)_{t+1}$ must price $r_{t+1}$ correctly conditional on $Z_t$, which results in

$$\phi_N = E_t[r_{t+1} m_R(\nu_t)_{t+1}] = \nu_t c_0^r Z_t + -\sigma_R \nu_t^{\prime} (\nu_t)_{t+1} c_0^r Z_t$$
$$+ (\nu_t^{\prime} (\nu_t)_{t+1} c_0^r Z_t)$$
$$- \nu_t^{\prime} (\nu_t)_{t+1} c_0^r Z_t = 0:$$

(39)
In case there is spanning this condition must again hold for every \( v_t \), implying

\[
(c_0^i - r_R^i \cdot k_R c_0^i)Z_t = 0 \quad \text{and} \quad (- r_R^i \cdot \frac{1}{k_R} \varphi^i_{k} \mid \varphi_{N}) = 0;
\]

(40)

It turns out that the regression framework that we used to test for spanning and intersection can be modified to test the restrictions in (39) and (40). Straightforward use of the algebra of partitioned matrices shows that in the regression

\[
r_{t+1} = cZ_t + dR_{t+1} + u_{t+1},
\]

(41)

with \( E[u_{t+1}Z_t] = 0 \), and \( E[u_{t+1}R_{t+1}] = 0 \), the OLS-estimates of \( c \) and \( d \) are consistent estimates of \((c_0^i - r_R^i \cdot \frac{1}{k_R} c_0^i)\) and \((- r_R^i \cdot \frac{1}{k_R} \varphi^i_{k} \mid \varphi_{N})\) respectively, which are the parameters of interest in the restrictions in (39) and (40) (see DeRoon, Nijman, and Werker (1998)). The hypotheses of intersection and spanning can therefore be based on the OLS-estimates of (41). The hypothesis that there is intersection for a given value of \( v_t \) and \( Z_t \) can be tested by testing the restrictions

\[
cZ_t v_t + (d \varphi^i_{k} \mid \varphi_{N}) = 0;
\]

(42)

and the hypothesis of spanning by testing the restrictions

\[
cZ_t = 0 \quad \text{and} \quad (d \varphi^i_{k} \mid \varphi_{N}) = 0;
\]

(43)

These restrictions are very similar to the restrictions implied by intersection and spanning in the unconditional case, except that the intercept \( \bar{\varphi} \) in (20) is replaced by \( cZ_t \).

It can easily be seen from (42) and (43) that the number of restrictions to be tested for intersection and spanning is the same as in the unconditional case, which makes this method of incorporating conditional information more parsimonious than using scaled returns. Note that the hypotheses underlying (42) and (43) are that there is intersection or spanning for a particular value of \( Z_t \), i.e., for a particular state of the economy. This has the additional advantage that the regression estimates of (41) make it possible to derive confidence intervals for the values of \( Z_t \) for which there can be intersection or spanning.

If the hypothesis of interest is whether there is spanning regardless of the state of the economy, the restrictions in (43) should hold for all values of \( z_t \), implying that each element of \( c \) should be equal to 0. In that case, with \( L \) instruments and \( N \) assets in \( r_{t+1} \), there are \( L \times N \) restrictions to be
tested, which, although smaller than the $2 \times L \times \mathbb{N}$ restrictions in (36), can be a large number. Also, as follows readily from (42) and (43), in this case the hypothesis of intersection and the hypothesis of spanning both imply the same restrictions. This latter result is due to the fact that the value of $v_t$ for which we test intersection is constant. Since the tangency point on the mean-variance frontier that corresponds to $v_t$ is a function of $Z_t$, the only way to have intersection irrespective of the specific value of $Z_t$ is to have spanning.

4.3 Regression coefficients linear in the conditional variables

An alternative way of incorporating conditional information in the regression framework is suggested by Shanken (1990) and Ferson and Schadt (1996) e.g., where the coefficients $\beta$ and $\gamma$ are assumed to be a linear function of the instruments. In the regression in (20), the $i$th row can be written as

$$r_{i,t+1} = \beta_i + \gamma_i R_{t+1} + \delta_{i,t+1};$$

Shanken (1990) simply assumes that

$$\beta_i = a_{i0} + z_t^\prime \beta_i;$$

$$\gamma_i = b_{i0} + z_t^\prime b_i;$$

where $z_t$ are now supposed to be $L$ demeaned variables. Here $a_{i0}$ is scalar, $a_i$ is an $L$-vector, $b_0$ is a $K$-row-vector, and $b_i$ is $L \times K$ matrix. Ferson and Schadt (1996) motivate (44) as a first order Taylor-series expansion for a general dependence of $\gamma$ on $Z_t = (1 z_t^0)^0$. Let $\text{Cov}[r_{t+1}; R_{t+1} j Z_t] = \Sigma_{RR}(Z_t)$, and $\text{Var}[R_{t+1} j Z_t] = \Sigma_{R}(Z_t)$, where $\Sigma(\cdot)$ indicates some functional form for the covariance matrix. Starting from (13) intersection for a given zero-beta rate $\gamma_t = 1$ conditional on $Z_t$ means

$$E[r_{t+1} j \gamma_t] = \gamma(Z_t) E[R_{t+1} j \gamma_t],$$

$$r_{t+1} j \gamma_t = \gamma(Z_t) (R_{t+1} j \gamma_t) + u_{t+1};$$

with $\gamma(Z_t) = \Sigma_{RR}(Z_t) \Sigma_{R}(Z_t)^{-1}, u_{t+1} \gamma (r_{t+1} j \gamma(Z_t) R_{t+1}) j (E[r_{t+1}] j \gamma(Z_t) E[R_{t+1}])$, and $E[u_{t+1} j Z_t] = 0$. Ferson and Schadt (1996) suggest a linear approximation of $\gamma(Z_t)$:

$$\gamma(Z_t) \approx b_0 + z_t^\prime b_1;$$
from which

\[ r_{i;t+1} = a_{i0} + z_{i0}a_{i1} + b_{00}R_{t+1} + (z_{i0}b_{10})R_{t+1} + \tilde{a}_{i+1}; \]

\[ a_{i0} = \hat{\gamma}_t \left( 1 + b_{00} \right) ; \]

\[ a_{i1} = \hat{\gamma}_t b_{10} \]

with \( u_{i:t+1} = u_{i:t+1} + \left( -\hat{\gamma}_t (Z_t) + b_{00} \right) (z_{i0}b_{10})\), for which it is assumed that \( E[u_{i:t+1} | Z_t] = 0 \). This yields precisely the regression in (20) where the regression parameters are linear in the instruments as assumed by Shanken (1990).

Intersection for a given value of \( \hat{\gamma}_t = 1 = v_t \) and \( z_t \) can now be tested by testing the restrictions that

\[ (a_{i0} + z_{i0}a_{i1}) + f(b_{00} + z_{i0}b_{10}) = 1 \]

As in the previous section, these restrictions have the additional advantage that statements can be made about in which state of the economy, (i.e., values of \( z_t \)) there is intersection. If there is intersection for all values of \( z_t \), this implies

\[ a_{i0} + (b_{00} + z_{i0}b_{10}) = 0 \]

\[ a_{i1} + b_{10} = 0 \]

The regression in (46) can also be motivated from the scaled returns in Section 4.1. Using the pricing kernel that is linear in \( R_{Z_t+1}^{\hat{\gamma}_t} \) and that is supposed to price the returns \( r_{Z_t+1} \) as well, the restrictions implied by intersection are very similar to the ones in (48). Thus, the use of managed returns is similar to the coefficients in the spanning regression being linear in the instruments.\(^3\)

Spanning for a given value of \( z_t \) is equivalent to

\[ a_{i0} + z_{i0}a_{i1} = 0 \]

\[ (b_{00} + z_{i0}b_{10}) = 1 \]

Again, for a specific value of \( z_t \), i.e., for specific economic conditions, these restrictions can easily be tested in the regression framework outlined above. If there is to be spanning under all economic conditions the restrictions are

\[ a_{i0} = 0 \]

\[ b_{00} = 1 \]

\[ a_{i1} = 0 \]

\[ b_{1} = 0 \]

\(^3\)We thank the referee for pointing this out to us.
If there are \( L \) instruments (including a constant) with \( K \) benchmark assets and \( N \) new assets, we now have \((K + 1)\) \(\leq N\) \(\leq L\) restrictions to test, which is even larger than with the scaled returns in Section 4.1. Also, the numbers of parameters to be estimated is \((K + 1)\) \(\leq N\) \(\leq L\). Thus, in terms of the number of parameters and the number of restrictions, this approach does not offer additional benefits over the use of scaled returns. However, this approach does have the benefit that it shows under what economic circumstances there may or may not be intersection or spanning.

Notice that this way of incorporating conditional information is very similar to the one suggested in the previous section. The restrictions on the regression parameters in (46) are analogous to the ones on the parameters in (41). The main difference arises because the slope coefficients for \( R_{t+1} \) also depend on the instruments, implying that the interaction term \( z_t R_{t+1} \) should also be included in the regression. It is easy to see that the approach in the previous section can be interpreted as a special case of the approach outlined here, where only the intercepts in (20) are a function of the instruments \( z_t \), whereas the slope coefficients are constant.

Summarizing, we have shown that a number of approaches is available to incorporate conditioning information in tests for intersection and spanning. Using either scaled returns or regression coefficients that are linear functions of the instruments, the regression approach outlined in Section 3 can easily be extended to test for intersection or spanning. The restrictions implied by the hypotheses of intersection and spanning are very similar to the case where there is no conditioning information (i.e., where the only instrument is a constant) and have very similar interpretations as well. Our methods focus on specific functional forms of incorporating conditioning information.

5 The relation between spanning tests, performance evaluation and optimal portfolio weights

So far the focus has been on the restrictions that are implied by the hypotheses of intersection and spanning on the distribution of \( R_{t+1} \) and \( r_{t+1} \) and on tests of these hypotheses. In this section the interest will be in the deviations from the restrictions. We will show that the test statistics and regression estimates have clear interpretations in terms of performance mea-
ures like Jensen's alpha and the Sharpe ratio as well as in terms of the new optimal portfolio weights. Since it is natural to think about these performance measures in terms of mean-variance efficient portfolios, most of the analysis in this section will be in terms of mean-variance frontiers rather than volatility bounds. Nonetheless, the duality between these two frontiers also holds for these performance measures. These interpretations of tests for mean-variance efficient, intersection, and spanning in terms of performance measures can also be found in Cochrane (1996), Dahlquist and Söderlind (1999), Gibbons, Ross and Shanken (1989), Jobson and Korkie (1982, 1984, 1989), and Kandel and Stambaugh (1989).

5.1 Performance measures

To set the stage, define the vector of Jensen's alphas, or Jensen performance measures, \( \tilde{\alpha}(\gamma) \), as the intercepts in a regression of the \( N \) excess returns \( (r_{t+1} - \gamma) \) on the excess returns of the \( K \) benchmark assets, \( (R_{t+1} - \gamma) \):

\[
r_{t+1} - \gamma = \tilde{\alpha}(\gamma) + (R_{t+1} - \gamma) + \varepsilon_{t+1};
\]

(50)

with \( E[\varepsilon_{t+1}] = E[(r_{t+1}R_{t+1})] = 0 \). Since it is not assumed that there exists a risk free asset, we define excess returns as the return on an asset or portfolio in excess of a given zero-beta rate \( \gamma \). Alternatively, when regressing \( r_{t+1} \) on \( R_{t+1} \) as in (20), it follows that Jensen's alpha is equal to

\[
\tilde{\alpha}(\gamma) = \tilde{\alpha} + (\tilde{\gamma}_k - \tilde{\gamma}_N); \quad (51)
\]

where \( \tilde{\alpha} = 1_r^\dagger \gamma_1_R \) and \( \tilde{\gamma} = \varepsilon \gamma, \tilde{\gamma}_R = \varepsilon \gamma, \tilde{\gamma}_R = \varepsilon \gamma \). Notice from this expression that the hypothesis that there is intersection for a given value of \( \gamma \) is equivalent to the hypothesis that the Jensen performance measure is zero, i.e., \( \tilde{\alpha}(\gamma) = 0 \). Similarly, the hypothesis of spanning is equivalent to the hypothesis that \( \tilde{\alpha}(\gamma) = 0 \), \( \tilde{\gamma} \). Recall from Section 3.3, that the regression in (50) produces the same intercept \( \tilde{\alpha}(\gamma) \) as a regression of \( r_{t+1} - \gamma \) on the excess return of a portfolio \( w_t^\gamma \) that is mean-variance efficient for \( R_{t+1} \) and that has \( \gamma \) as its zero beta rate, i.e.,

\[
r_{t+1} - \gamma = \tilde{\alpha}(\gamma) + (R_{t+1} - \gamma) + \varepsilon_{t+1};
\]

Following Jensen (1968), it is common in the literature to define Jensen's alpha as the intercept of a regression of \( r_{t+1} \) in excess of the risk free rate on
the return of the market portfolio in excess of the risk free rate. The definition in (50) is more general and has this more traditional definition as a special case if there exists a risk free asset \( \hat{R}_t \) and if the market portfolio is mean-variance efficient \( R_m^{t+1} = R_m^{t+1} \). The Jensen measure in (50) is also referred to as the generalized Jensen measure. Given the minimum variance stochastic discount factor \( m_R(v)_{t+1} \) as defined in (4) and (5), it can easily be seen that the generalized Jensen measure is also equal to \( \mu(v) \equiv v \) as defined in (28). This is also discussed in Cochrane (1996) and in Dahlquist and Söderlind (1999).

The Sharpe ratio of a portfolio with return \( R_p^{t+1} \) is defined as the expected excess portfolio return, divided by the standard deviation of portfolio return,

\[
Sh(R_p^{t+1}; \hat{\cdot}) = \frac{E[R_p^{t+1}] - \hat{\cdot}}{\sqrt{\text{Var}(R_p^{t+1})}}.
\]

By definition, for a given expected portfolio return, or for a given standard deviation of portfolio return, the maximum attainable (absolute) Sharpe ratio is the Sharpe ratio of the minimum-variance efficient portfolio. For a minimum-variance efficient portfolio \( w_R^{\ast} \) of the \( K \) assets \( R_t^{t+1} \) with zero-beta rate \( \hat{\cdot} \), the Sharpe ratio is equal to the slope of the line tangent to the frontier originating at \((0; \hat{\cdot})\) in mean-standard deviation space, and is denoted by \( \mu_R(\hat{\cdot}) \):

\[
\mu_R(\hat{\cdot}) = \frac{E[R_p^{t+1}] - \hat{\cdot}}{\sqrt{\text{Var}(R_p^{t+1})}};
\]

(52)

where \( R_p^{t+1} \equiv w_R^{\ast}R_t^{t+1} \).

Although both Jensen's alpha and the Sharpe ratio are used as performance measures, there is an important difference between the two. Whereas the Sharpe ratio is defined in terms of the characteristics of one portfolio (the expected excess portfolio return and its standard deviation), Jensen's alpha is defined in terms of one asset or portfolio relative to another. Sharpe ratios answer the question whether one portfolio is to be preferred over another, whereas Jensen's alpha answers the question whether investors can improve the efficiency of their portfolio by investing in the new asset. However, there is a close relation between the two measures, in that Jensen's alphas together with the covariance matrix of the error terms \( u_t^{t+1} \) in (20) (and (50)) determine the potential improvement in the maximum attainable Sharpe ratio from adding the new assets \( r_t^{t+1} \). Recall from Section 2.2 that we defined the variables \( A \equiv \Sigma_1 \), \( B \equiv \Sigma_1 \), and \( C \equiv \Sigma_1 \). For the set \( R_t^{t+1} \) these
variables will be denoted as $A_R$, $B_R$, and $C_R$, whereas the absence of subscripts implies that these variables refer to the larger set $(R_{t+1}; r_{t+1})$. Using partitioned inverses, notice that

\[
\begin{align*}
\Sigma^{-1} & = \begin{bmatrix}
\mathbf{A} & \mathbf{B} \\
\mathbf{B}^T & \mathbf{C}
\end{bmatrix}^{-1} \\
\mathbf{A} & = \mathbf{A}_R + \mathbf{B}_R \mathbf{r}_R^{-1} \mathbf{C}_R + \mathbf{B}_R \mathbf{r}_R^{-1} \mathbf{C}_R + \mathbf{B}_R \mathbf{r}_R^{-1} \mathbf{C}_R
\end{align*}
\]

From this, it follows that

\[
\begin{align*}
\mu_R(\gamma) & = (C_R \mathbf{i} + 2B_R \gamma + A_R \gamma^2)^{1/2} \\
\mu_R^2 & = (C_R \mathbf{i} + 2B_R \gamma + A_R \gamma^2)
\end{align*}
\]

A similar expression holds of course for $\mu(\gamma)$, the maximum attainable Sharpe ratio of the larger set $(R_{t+1}; r_{t+1})$. Using (54) and (55), we derive

\[
\begin{align*}
\mu(\gamma) & = C_R \mathbf{i} + 2B_R \gamma + A_R \gamma^2 \\
\mu_R^2 & = (C_R \mathbf{i} + 2B_R \gamma + A_R \gamma^2)
\end{align*}
\]

Thus, the change in maximum attainable squared Sharpe ratios equals the inner product of the vector of Jensen's alphas, $\mathbf{\beta}(\gamma)$, weighted by the inverse of the covariance matrix of $\mu_R(\gamma)$. If there is only one new asset, $N = 1$, the term $\mathbf{\beta}(\gamma) = \theta^2$ is known as the adjusted Jensen measure or the appraisal ratio (Treynor and Black (1973)). Notice once more that $\mu_R(\gamma)$ and $\mu(\gamma)$
characterize portfolios of $R_{t+1}$ and $(R_{t+1}^0; r_{t+1}^0)$, respectively, whereas $\tilde{\alpha}(\gamma)$ and $\tilde{\gamma}$ follow from a regression of $r_{t+1}$ on $R_{t+1}$, and measure the performance of $r_{t+1}$ relative to $R_{t+1}$. Stated differently, whereas Sharpe ratios can be used to compare the performance of different portfolios, Jensen’s alpha gives the potential improvement in performance when the additional assets are included in the portfolio. The hypotheses of intersection and spanning imply that Jensen’s alpha, $\tilde{\alpha}(\gamma)$, is zero for one or for all values of $\gamma$ respectively. Therefore, if there is intersection (spanning) then there is no improvement in the Sharpe measure possible by including the additional assets $r_{t+1}$ in the investor’s portfolio.

Cochrane and Saá-Requejo (1996) show how a bound on the maximum Sharpe ratio can be used to price new assets in incomplete markets, which is referred to as “good deal” pricing. In the context of (57) this essentially comes down to putting a bound on the maximum appraisal ratios of the new asset. This kind of analysis is extended by Bernardo and Ledoit (1996), who introduce the gain-loss ratio as an alternative performance measure by which new assets can be priced if restrictions on the maximum gain-loss ratio are imposed. This is similar to a bound on the maximum Sharpe ratio as suggested by Cochrane and Saá-Requejo (1996), but the approach in Bernardo and Ledoit (1996) is more general and allows for non-mean variance utility functions as well.

5.2 Changes in optimal portfolio weights

The performance measures and the intersection regressions discussed above can also be used to infer the changes in optimal portfolio holdings when adding the assets $r_{t+1}$. In this section we will show that given the initial mean-variance efficient portfolio of the benchmark assets and the OLS-estimates of the regression parameters in (20), it is straightforward to determine the new optimal portfolio weights. Some of the results presented in this section are also presented in Stevens (1998). In order to derive the optimal portfolio weights from the regression results, consider the mean-variance efficient portfolio for the extended set $(R_{t+1}; r_{t+1})$ for a given value of $\gamma$:

$$w^* = \omega^* \tilde{\gamma} \tilde{\gamma}^{-1} \{ 1 \}$$

Substituting the partitioned inverse as given in (53) in the expression for $w^*$ gives that the optimal portfolio weights for the new assets, $w^*_r$, can be written
Thus, the optimal portfolio weights $w^a$ are determined by the vector of Jensen's alphas and the covariance matrix of the residuals of the OLS-regression of $r_{t+1}$ on $R_{t+1}$. This result is simply a generalization of the well known result in Treynor and Black (1973) regarding the appraisal ratio. The difference with Treynor and Black is that these authors assume that the error terms "$i,t+1$ for different securities are uncorrelated, i.e., they assume the diagonal model (Sharpe (1963)), whereas the result in (58) allows for any correlation structure between the securities.

In deriving the new optimal portfolio weights, a problem in (58) is that the coefficient of risk aversion $\alpha$ is present. Notice that this is a different coefficient than the one that appears in the optimal portfolio $e_{w}^{R}$ of the smaller set $R_{t+1}$:

$$w^a_R = e_R^{1}S_R^{1/2}(\mu_{R} + \Lambda J(\gamma)),$$

where we now also add a ~ to indicate that a variable refers to the set of benchmark assets $R_{t+1}$ only. It is only the zero-beta return $\gamma$ that is the same in both problems, since we test whether there is intersection for a fixed value of $\gamma$. Similarly, the expected returns on the portfolios $w^a_R$ and $w^a$ are different, and we indicate these with $m_R$ and $m$ respectively, i.e., $m_R w^a_R R$ and $m w^a$. In order to substitute out the risk aversion parameter $\alpha$, note that

$$\alpha = B i A = B R j A_R + @ (\gamma)S j^{1/2}(\mu_N i E_k)$$

and that

$$e_R = \frac{m_R i}{w_R^{aR} E_R^{1/2} w_R^{a'}} = \frac{m_R^{2}}{m_R i},$$

As an aside, in terms of volatility bounds, notice that $w_{R}^{a'} = m^{'} (1=')$, i.e., the elements of $' (v)$ in (5) that correspond to $r_{t+1}$. Thus if we want to know the minimum variance stochastic discount factor from $(R_{t+1};r_{t+1})$, rather than from $R_{t+1}$, the projection coefficients corresponding to the additional assets $r_{t+1}$ are given by $i S_{j}^{1/2} @ (\gamma)'$. 

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Using these latter two expressions, the optimal portfolio weights \( w_i^* \) can be expressed as

\[
w_i^* = \frac{\hat{A}}{\mu_r^2 + (\hat{m}_r i \cdot \hat{\beta} i)\hat{\sigma} i \cdot \hat{\beta} i + \hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{\phi} i \cdot \hat{\phi} i \cdot \hat{f}_i}{\hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{f}_i}{\hat{f}_i}.
\]

The advantage of (59) is that it contains only variables that either result from the initial optimal portfolio \( w_R^* \), or from a regression of \( r_{t+1} \) on \( R_{t+1} \).

Along the same lines it is straightforward to show that the new optimal weights \( w_R^* \) are given by

\[
w_R^* = \frac{\mu_r^2}{\mu_r^2 + (\hat{m}_r i \cdot \hat{\beta} i)\hat{\sigma} i \cdot \hat{\beta} i + \hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{\phi} i \cdot \hat{\phi} i \cdot \hat{f}_i}{\hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{f}_i}{\hat{f}_i} \frac{\hat{w}_R^*}{\hat{w}_R^*} - \hat{w}_R^*.
\]

Again, this expression only depends on characteristics of the old portfolio, \( w_R^* \), and the regression output. Therefore, given the initial mean-variance efficient portfolio \( w_R^* \) of the benchmark assets and the OLS-estimates of the regression in (20), equations (59) and (60) answer the question how to adjust the portfolio in order to obtain the new mean-variance efficient portfolio \( w^* \).

In order to give an interpretation of the new portfolio weights in (59) and (60), it is useful to rewrite them in the following way:

\[
w_i^* = \frac{m_i \cdot \hat{\beta} i}{\mu_r^2} \frac{\hat{\phi} i \cdot \hat{\phi} i \cdot \hat{f}_i}{\hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{f}_i}{\hat{f}_i};
\]

and

\[
w_R^* = \frac{\mu_r^2}{\mu_r^2 + (\hat{m}_r i \cdot \hat{\beta} i)\hat{\sigma} i \cdot \hat{\beta} i + \hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{\phi} i \cdot \hat{\phi} i \cdot \hat{f}_i}{\hat{\phi} i \cdot \hat{\phi} i} \frac{\hat{f}_i}{\hat{f}_i} \frac{\hat{w}_R^*}{\hat{w}_R^*} - \hat{w}_R^*.
\]

If there is only one new asset, i.e., \( N = 1 \), Equation (61) reveals all of the signs that determine the sign of the new portfolio weight \( w_i^* \) (given that \( m_i \cdot \hat{\beta} i > 0 \)): Jensen's alpha is positive (negative) if Jensen's alpha is positive (negative) the investor can improve the performance of his portfolio by taking long (short) positions in the new asset. When there is more than one new asset, the sign of the portfolio weights is not only determined by the sign of Jensen's alpha, but also by the inverse of the covariance matrix of \( r_{t+1} \). If the mean-variance frontier is not strongly affected by the introduction of the new assets, then

\[\text{Here we use the fact that } \mu_r^2 = \hat{m}_r i \cdot \hat{\beta} i = A_R i \cdot \hat{\beta} R, \text{ and that } A_R i \cdot \hat{\beta} R + \hat{\phi} i \cdot \hat{\phi} i = A i \cdot \hat{\phi} R.\]
\((\mu_R(\cdot)^2 = \mu(\cdot)^2)(m_1 \cdot \cdot = f_{R_1, \cdot \cdot})^{1/41}\), and the coefficients show which of the old assets are replaced by the new assets.

Finally, notice that we did not consider a risk free asset. The portfolio weights given above are for the tangency portfolio when the zero-beta rate is \(\cdot\). If a risk free asset is available, we can replace \(\cdot\) with \(R_f\) in (61) and (62) and these equations still give the portfolio weights for the tangency portfolio. The new tangency portfolio has an expected return equal to \(m\), whereas the old tangency portfolio has an expected return \(f_{R_1, \cdot \cdot}\). Notice though, that in case a risk free asset is available it is easy to shift funds between the tangency portfolio and the risk free asset and to let the expected portfolio return vary. For practical purposes, the interest may be in the new portfolio \(w^\alpha\) that has the same expected return as the old portfolio. Given that there is a risk free asset available, this is easily achieved by letting \(m_1 \cdot R_f = f_{R_1, \cdot \cdot} R_f\). In this case Equations (61) and (62) simplify to

\[
\begin{align*}
  w^\alpha_r &= \frac{m_1 \cdot R_f}{\mu^2} - \xi \cdot ^1 \cdot \beta_j \\
  w^\alpha_R &= \frac{\mu^2}{\mu^2} w_{R_1, \cdot \cdot} \cdot ^{-0} w^\alpha_r;
\end{align*}
\]

Notice that here it is not necessarily the case that the weights in \(w^\alpha_r\) and \(w^\alpha_R\) sum to one. The investor will have to borrow or lend a fraction \((1 - \xi \cdot \beta_j \cdot ^{1} \cdot \beta_j)\) to achieve an expected portfolio return equal to \(m\).

### 5.3 Interpretation of spanning and intersection tests in terms of performance measures

Finally, we want to relate the Wald test-statistics presented in Section 3 to the performance measures discussed above. It will be shown that these test-statistics can be expressed as changes in maximum Sharpe ratios of \(R_{t+1}\) and \((R_{t+1}; r_{t+1})\) respectively. Therefore, they have a clear economic interpretation. In order to interpret the test-statistics for intersection and spanning in terms of performance measures, recall the basic regression model in (20):

\[
r_{t+1} = \delta + (\cdot \cdot R_{t+1} + f_{t+1};
\]

where intersection for a given value of \(\cdot\) means that

\[
\delta (\cdot) = \delta + (\cdot \cdot w_{R_1, \cdot \cdot} \cdot ^{-0} \cdot W_{R_1, \cdot \cdot} \cdot ^{0} \cdot W_{R_1, \cdot \cdot});
\]

\[34\]
Thus, the restrictions on the regression coefficients that are imposed by the hypothesis of intersection have a natural interpretation in terms of Jensen's alphas, and - as noted before - testing whether there is intersection for \( \hat{\beta} \), is equivalent to testing whether Jensen's alpha is zero. Testing for spanning is of course equivalent to testing whether the Jensen's alphas are zero for all values of \( \hat{\beta} \).

It can be shown that the test statistics for intersection and spanning, \( \chi^2_{int} \) and \( \chi^2_{span} \), presented in Section 3.4, can also be interpreted in terms of Jensen's alphas and Sharpe ratios. To see this, start again from the specification of the regression equation in (23):

\[
\begin{align*}
    r_{t+1} &= \mathbf{1}_N - 1 \mathbf{R}_t^0 \mathbf{b} + \eta_{t+1}
\end{align*}
\]

Note that (using partitioned inverses) the asymptotic covariance matrix of the OLS-estimates of \( \mathbf{b} \) in (23) is given by

\[
\begin{align*}
    \text{Var}[\mathbf{b}] &= \mathbf{H} \begin{bmatrix} \hat{\mu} & \hat{\beta} \end{bmatrix} (\hat{\beta}^2) \begin{bmatrix} \hat{\mu} & \hat{\beta} \end{bmatrix}
\end{align*}
\]

Straightforward algebra shows that premultiplying (65) with \( \mathbf{H}(\hat{\beta}) \) and postmultiplying with \( \mathbf{H}(\hat{\beta})^\top \) as defined in (25), yields

\[
\begin{align*}
    \text{Var}[\hat{\beta}] &= \mathbf{H}(\hat{\beta}) \begin{bmatrix} \hat{\mu} & \hat{\beta} \end{bmatrix} (\hat{\beta}^2) \begin{bmatrix} \hat{\mu} & \hat{\beta} \end{bmatrix} \mathbf{H}(\hat{\beta})^\top \\
    &= \mathbf{H}(\hat{\beta}) \begin{bmatrix} \hat{\mu} \hat{\mu} & \hat{\mu} \hat{\beta} + \hat{\beta} \hat{\mu} \\
    &\hat{\beta} \hat{\beta} \end{bmatrix} \begin{bmatrix} \hat{\mu} & \hat{\beta} \end{bmatrix} \mathbf{H}(\hat{\beta})^\top
\end{align*}
\]

where the Sharpe ratio \( \hat{\mu}(\cdot) \) was defined in (56). Since from the analysis above we know that the term \( h(\cdot)_{int} \) as defined in (25) equals \( \hat{\beta}(\cdot) \), (57) can be used to rewrite the test statistic for intersection, \( \chi^2_{int} \), as

\[
\chi^2_{int} = \begin{bmatrix} \hat{\beta}(\cdot) \hat{\mu}(\cdot) \end{bmatrix} \begin{bmatrix} \hat{\mu}(\cdot) & \hat{\beta}(\cdot) \end{bmatrix} \begin{bmatrix} \hat{\mu}(\cdot) & \hat{\beta}(\cdot) \end{bmatrix} \begin{bmatrix} \hat{\beta}(\cdot) \hat{\mu}(\cdot) \end{bmatrix}
\]

where \( \hat{\beta}(\cdot) \) and \( \hat{\beta}(\cdot) \) are the sample Sharpe ratios and Jensen's alpha respectively. Equation (67) is a well known result from, e.g., Jobson and Korkie (1982) and Gibbons, Ross and Shanken (1989). It clearly shows that the Wald test statistic for intersection can easily be interpreted as the percentage increase in squared Sharpe ratios scaled by the sample size. Under the null-hypothesis that there is intersection, \( \mu(\cdot) = \mu(\hat{\beta}) \) and the increase
of the sample Sharpe ratios scaled by the sample size $T$ (as in (67)) will asymptotically have a $\chi^2_{(N)}$-distribution.\footnote{Gibbons, Ross, and Shanken (1989) study the small sample properties of this test statistic in case there is a risk free asset, as well as the distribution under the alternative hypothesis. Kandel and Stambaugh (1987) and Shanken (1987) extend their results to the case where the mean-variance efficient benchmark portfolio (or intersection portfolio) can not be observed but has a given correlation with the observed proxy portfolio.}

MacKinlay (1995) uses a similar interpretation of the Wald test-statistic in case returns are normally distributed together with (57) to distinguish between risk-based alternatives for the CAPM and nonrisk-based alternatives. His analysis suggests that for reasonable values of the maximum attainable Sharpe ratios a multifactor model like the one proposed by Fama and French (1996) can not explain the deviations from the CAPM that are found in the cross section of asset returns.

For the spanning test-statistic, a similar interpretation can be given. Let $\gamma_R^0$ denote the expected return on the global minimum variance portfolio of $R_{t+1}$, i.e., $\gamma_R^0 = B^0_R \gamma_R$, and let the variance of this portfolio be given by $(\gamma_R^0)^2$. Similarly, let $(\gamma^0)^2$ be the global minimum variance of $(R_{t+1}, r_{t+1})$. It is shown in Appendix B that the Wald test-statistic for spanning, $\gamma_{span}^w$, can be written as

$$\gamma_{span}^w = T \left( \frac{1 + B^0_R (\gamma_R^0)^2}{1 + B_R ((\gamma^0)^2)} \right) + T \left( \frac{1}{(\gamma^0)^2} \right)$$

This shows that the spanning test-statistic consists of two parts. The first part is similar to the test-statistic for intersection in (67) and is determined by a change in Sharpe ratios. The Sharpe ratios in (68) are for a zero-beta rate equal to the (in-sample) expected return on the global minimum variance portfolio however, and therefore are the slopes of the asymptotes of the mean-variance frontier. Notice that the slope of the upper limb of the frontier is simply the negative of the slope of the lower limb of the frontier, and therefore, the squared Sharpe ratios for those two extremes are the same. The first term of the spanning test-statistic in a sense measures whether there is intersection at the most extreme points of the frontier (i.e., whether there is a limiting form of intersection if we go sufficiently far up or down the frontier). The second term of the statistic in (68) is determined by the change in the global minimum variance of the portfolios, and measures whether the point most to the left on the frontier changes or not. Put differently, the first term measures whether there is intersection for a mean-variance investor.
with a very small risk aversion \( (\gamma = 0) \), while the second term measures whether there is intersection for a mean-variance investor with a very high risk aversion \( (\gamma \neq 1) \). Note that in the second term the old global minimum variance appears in the numerator and the new global minimum variance in the denominator, since this variance can only decrease as assets are added to the portfolio. Therefore, both terms in (68) are always larger than or equal to one. Jobson and Korkie (1989) derive a similar result for a likelihood ratio test for spanning.

6 Specification error bounds and intersection

As in the previous section, in this section the focus will be on deviations from intersection rather than on intersection itself. In a recent paper Hansen and Jagannathan (1997) analyze specification errors in stochastic discount factor models which, in some special cases, can be interpreted as deviations from intersection. They derive bounds on the magnitude of these specification errors.

Recall from the discussion in Section 2.1 that each asset pricing model assigns a particular function to the pricing kernel \( M_{t+1} \). Hansen and Jagannathan (1997) note that the pricing kernels implied by most asset pricing models do not yield correct asset prices, either because the asset pricing model can only be viewed as an approximation, or because of measurement error. Measurement errors are for instance often considered to be an important problem in measuring consumption and testing consumption based asset pricing models. Therefore, the pricing kernel implied by an asset pricing model will in general only serve as a proxy stochastic discount factor, that will not yield the correct prices or expected payoffs of the assets under consideration. In a related paper Balduzzi and Robotti (2000) focus on the estimation of risk premia as a separate problem from the testing of asset pricing models. They estimate risk premia by looking at the prices assigned by the minimum variance kernel to risk variables, or by the prices of hedge portfolios that are the linear projections of risk variables on asset returns.

The interest of Hansen and Jagannathan is in the least squares distance between a proxy stochastic discount factor and the set of valid stochastic discount factors. They derive a lower bound on this distance, the specification error bound, as a measure of how well the model performs. These specification error bounds will be derived formally below and it will also be shown
that these bounds have a clear economic interpretation in terms of maximum pricing errors or maximum expected payoff errors implied by the asset pricing model. Hansen, Heaton, and Luttmer (1995) derive the limiting distribution for the corresponding estimator of the specification error bounds.

It turns out that if we take the minimum variance stochastic discount factor for the subset \( R_{t+1} \) as a proxy stochastic discount factor for the larger set of assets \( (R_{t+1}; r_{t+1}) \), we can interpret the specification error bounds in terms of mean-variance intersection and the performance measures discussed in the previous section. In particular, provided that both the proxy stochastic discount factor and the discount factors that price \( R_{t+1} \) and \( r_{t+1} \) correctly have the same expectation \( v \), the squared specification error bound scaled by \( v \) turns out to be equal to the difference between the maximum squared Sharpe ratio implied by the set \( R_{t+1} \) and the maximum squared Sharpe ratio implied by \( (R_{t+1}; r_{t+1}) \). This also allows us to interpret the specification errors in terms of mean-variance portfolio choice again. Given that a mean-variance investor is aware of the fact that a portfolio chosen from the subset \( R_{t+1} \) is suboptimal relative to a portfolio chosen from the larger set \( (R_{t+1}; r_{t+1}) \), the specification error bound gives an estimate of the extent to which the portfolio is suboptimal in terms of Sharpe ratios.

### 6.1 Specification error bounds

As noted above, in Hansen and Jagannathan (1997) the interest is in proxy stochastic discount factors, denoted by \( y_{t+1} \), that assign approximate prices to portfolio payoffs. For instance, the CAPM implies that the proxy is of the form \( a + br_{t+1}^{m} \), with \( r_{t+1}^{m} \) the return on the market portfolio. As before, let \( R_{t+1}^{p} \) be the return on some portfolio, not necessarily mean-variance efficient, such that \( w_{p}^{\alpha}K_{t} = 1 \). The expected price assigned to such a portfolio by a proxy stochastic discount factor will be denoted by \( \frac{1}{2}(R_{t+1}^{p}) \):

\[
E[y_{t+1}R_{t+1}^{p}] = \frac{1}{2}(R_{t+1}^{p}) \quad (69)
\]

Of course, valid stochastic discount factors \( M_{t+1} \) would assign a price \( \frac{1}{2}(R_{t+1}^{p}) = 1 \) to any portfolio \( w^{p} \) that satisfies \( w_{p}^{\alpha}K_{t} = 1 \). Because the proxy \( y_{t+1} \) may be derived from an asset pricing model that is strictly speaking not valid, or because the proxy may be measured with error, the prices assigned by the proxy, \( \frac{1}{2}(R_{t+1}^{p}) \), will in general not be equal to one. We only consider payoffs that are returns, i.e., payoffs with (correct) prices equal to one. Hansen and
Jagannathan (1997) take more general payoffs $x_{t+1}$ with current prices $q_t$. Given that we want to establish the relation between specification errors and mean-variance intersection, the use of returns is not very restrictive however. Moreover, the results derived below can easily be adjusted to the results of Hansen and Jagannathan along the lines of Section 4.1, where we incorporated conditioning information by allowing for payoffs $z_t - R_{t+1}$ with current prices $q_t$.

A second way in which the results here are somewhat more restrictive than the ones in Hansen and Jagannathan (1997) is that we will always consider valid stochastic discount factors $M(v)_{t+1}$ that have the same expectation as the proxy $y_{t+1}$, i.e., $v = E[y_{t+1}]$. This may be considered as restrictive, since this assumption in fact requires that the proxy assigns the correct price to the risk free payoff, if it exists. Once more, given that the interest here is in the relation with mean-variance intersection in the absence of a risk free asset, and given that we always defined intersection for a known value of $v$, this is not restrictive for our purposes.

The problem addressed in Hansen and Jagannathan (1997) is to derive a lower bound $\pm$ on the distance between $y_{t+1}$ and the set of stochastic discount factors that price $R_{t+1}$ correctly, which we denote as $M$:

$$\pm = \min_{fM_R(v)_{t+1}} k y_{t+1} + M_R(v)_{t+1} k;$$

(70)

where $k x_{t+1} k = E[x_{t+1}^2]$. Because $y_{t+1}$ and $M_R(v)_{t+1}$ have the same expectation, the distance between $y_{t+1}$ and $M_R(v)_{t+1}$ in (70) is equal to the standard deviation of $y_{t+1} + M_R(v)_{t+1}$, i.e., $k y_{t+1} + M_R(v)_{t+1} k = \frac{1}{2} (y_{t+1} + M_R(v)_{t+1})$. We will denote the stochastic discount factor that solves (70) by $m_R(v)_{t+1}$. Thus, $m_R(v)_{t+1}$ is the stochastic discount factor that prices $R_{t+1}$ correctly and that is closest to $y_{t+1}$ in a least squares sense.

To solve the problem in (70), consider the least squares projections of $y_{t+1}$ and $M_R(v)_{t+1}$ on $R_{t+1}$ and a constant:

$$\begin{align*}
y_{t+1} &= \text{Proj}(y_{t+1} j R_{t+1}) = v + 3(v) q R_{t+1} \quad \text{and} \\
m_R(v)_{t+1} &= \text{Proj}(M_R(v)_{t+1} j R_{t+1}) = v' + (v) q R_{t+1} \quad \text{and}
\end{align*}
$$

(71)

and

$$\begin{align*}
m_R(v)_{t+1} &= \text{Proj}(M_R(v)_{t+1} j R_{t+1}) = v + 3(v) q R_{t+1} \quad \text{and} \\
M_R(v)_{t+1} &= m_R(v)_{t+1} + w_{t+1};
\end{align*}
$$

(72)

39
where \( m_R(v)_{t+1} \) is the minimum variance stochastic discount factor derived in Section 2.1, and \( \hat{v} \) is defined in (5). The projection coefficients in (71) are given by \( \hat{y}_{t+1} = \hat{m}_R(v)_{t+1} k^2 \). Noting that \( k y_{t+1} \cdot m_R(v)_{t+1} k^2 = Var[y_{t+1} \cdot m_R(v)_{t+1}] \), it easily follows that

\[
Var[y_{t+1} \cdot m_R(v)_{t+1}] = Var[y_{t+1} \cdot m_R(v)_{t+1}] + Var[u_{t+1} \cdot w_{t+1}]
\]

Because \( y_{t+1} \cdot m_R(v)_{t+1} = y_{t+1} \cdot (m_R(v)_{t+1} + u_{t+1}) \) and \( u_{t+1} \) is orthogonal to \( R_{t+1} \), this lower bound on the variance of \( y_{t+1} \cdot m_R(v)_{t+1} \) is attainable for the stochastic discount factor

\[
f_{R_{t+1}} = m_R(v)_{t+1} + u_{t+1}; \tag{73}
\]

and we have that

\[
\pm^2 = Var[y_{t+1} \cdot m_R(v)_{t+1}]; \tag{74}
\]

A more detailed characterization of \( f_{R_{t+1}} \) and \( \pm \) will be given in the following section. For this moment, note that subtracting the variable \( y_{t+1} \cdot m_R(v)_{t+1} \) from the proxy \( y_{t+1} \cdot f_{R_{t+1}} \) yields a valid stochastic discount factor. Therefore, as noted by Hansen and Jagannathan (1997), \( y_{t+1} \cdot f_{R_{t+1}} \) is the smallest adjustment in a least squares sense that is necessary to make \( y_{t+1} \) a valid stochastic discount factor, and \( \pm \) is a measure of the magnitude of this adjustment.

Hansen and Jagannathan also show that \( \pm \) can be interpreted as a maximum pricing error. In order to do so, let \( R_{t+1} \) denote a position in \( R_{t+1} \) that does not necessarily satisfy the requirement \( \|\eta_k\| = 1 \), i.e., \( \eta \) is in general not a portfolio. Denote the payoff of such a position as \( R(!)_{t+1} = \|\eta\|_{t+1} \) and note that the correct price of such a position is

\[
E[\|\eta\|_{t+1} \cdot m_R(v)] = \frac{1}{2} \langle R(!)_{t+1} \rangle = \|\eta\|_{t+1},
\]

whereas the price assigned by the proxy \( y_{t+1} \) is \( \frac{1}{2} \langle R(!)_{t+1} \rangle \). The pricing error of the proxy \( y_{t+1} \) is therefore \( \frac{1}{2} \langle R(!)_{t+1} \rangle \cdot \frac{1}{2} \langle R(!)_{t+1} \rangle \), and Hansen and Jagannathan show that \( \pm \) provides an upper bound on the absolute value of this pricing error, for positions that have a unit norm:

\[
\pm = \max_{R(!)_{t+1} \cdot \eta \approx R(!)_{t+1}} \frac{1}{2} \langle R(!)_{t+1} \rangle \cdot \frac{1}{2} \langle R(!)_{t+1} \rangle.
\]
Thus, by looking at a particular class of positions, i.e., positions with a unit norm, \( \pm \) can be interpreted as the maximum pricing error assigned by the proxy to the payoffs of those unit norm positions.

A more intuitive interpretation can be given if we consider errors in expected payoffs, or expected returns, rather than pricing errors. Recall that a valid stochastic discount factor assigns the correct expected return to a one-dollar investment in portfolio \( w^p \) (for which, by definition, \( w^p \| = 1 \)) which, using equation (3), can be written as

\[
E[R^p_{t+1}] = \frac{1}{v} i \frac{\text{Cov}[M_R(v)_{t+1}; R^p_{t+1}]}{v},
\]

i.e., as one over the expectation of the pricing kernel, which equals the risk free rate if it exists, plus a risk term which is determined by the covariance of the portfolio return and the pricing kernel. Observe that use of the proxy, that also has expectation \( v \), would give an approximate expected return \( E^a[R^p_{t+1}] \) for a one-dollar investment in \( w^p \) that in general differs from \( E[R^p_{t+1}] \), because the covariance of the proxy with the portfolio return will be different from the covariance of a valid stochastic discount factor with the portfolio return, i.e.:

\[
E^a[R^p_{t+1}] = \frac{1}{v} i \frac{\text{Cov}[y_{t+1}; R^p_{t+1}]}{v}.
\]

From these relations we define the expected return error

\[
E^a[R^p_{t+1}] - E[R^p_{t+1}] = \frac{\text{Cov}[M_R(v)_{t+1} i y_{t+1}; R^p_{t+1}]}{v};
\]

for which the Cauchy-Schwarz inequality implies that

\[
j E^a[R^p_{t+1}] i E[R^p_{t+1}] = \frac{\text{Cov}[M_R(v)_{t+1} i y_{t+1}; R^p_{t+1}]}{v};
\]

Since this inequality holds for all valid stochastic discount factors \( M_R(v)_{t+1} \), it also holds for the stochastic discount factor that solves (70), \( f_{M_R(v)_{t+1}} \), implying

\[
j E^a[R^p_{t+1}] i E[R^p_{t+1}] = \frac{\text{Cov}[M_R(v)_{t+1} i y_{t+1}; R^p_{t+1}]}{v};
\]

Since for a given value of \( v \), the Sharpe ratio is defined as \( \text{Sh}(R^p_{t+1}) = \frac{E[R^p_{t+1}]}{\text{Var}(R^p_{t+1})} \), and the approximate Sharpe ratio, i.e., the Sharpe
ratio according to the proxy $y_{t+1}$, as $\text{Sh}^\alpha(R^P_{t+1})$ \quad (E^\alpha[R^P_{t+1} | 1-v)=\gamma(R^P_{t+1})$, this can be rewritten as

$$j \text{Sh}^\alpha(R^P_{t+1}) \text{Sh}(R^P_{t+1}) j \pm \frac{\sigma}{v}$$  \hspace{1cm} (76)

Thus, using errors in expected returns rather than errors in assigned prices, the specification error bound $\pm$ scaled by the expectation of the proxy has a very clear interpretation in terms of Sharpe ratios. For any portfolio $w^P$ formed from the assets in $R_{t+1}$, the absolute difference between the approximate Sharpe ratio assigned to the portfolio returns by $y_{t+1}$ and the actual Sharpe ratio of the portfolio can never exceed the scaled specification error bound $\pm v$. This interpretation is also somewhat easier than the one given for the expected payoffs error in Hansen and Jagannathan (1997), where they focus on the maximum error in expected payoffs for positions with unit standard deviation.

### 6.2 The relation between specification error bounds and intersection

The purpose of this section is to show that there is a close relation between intersection and a special case of the specification error bounds. In particular, if the interest is in stochastic discount factors that price the returns $(R_{t+1}; r_{t+1})$ correctly and we choose for the proxy $y_{t+1}$ the minimum variance stochastic discount factor based on the subset $R_{t+1}$, $m_R(v)_{t+1}$, the specification error bound can simply be expressed as a deviation from intersection, as was the case with the performance measures discussed in Section 5. To show this, let us first give a more precise characterization of $f_{m_R(v)_{t+1}}$ and $\pm$ than given in (73) and (74).

Recall that $f_{m_R(v)_{t+1}}$ is given by $m_R(v)_{t+1} + u_{t+1}$, where $u_{t+1} = y_{t+1} - v_{t+1}$. Using (72) and (71), this implies for $f_{m_R(v)_{t+1}}$:

$$f_{m_R(v)_{t+1}} = v + \gamma(R_{t+1}^{1_R} + y_{t+1} - f_{m_R(v)_{t+1}} g$$

$$= y_{t+1} + (\gamma(R_{t+1}^{1_R}) + f_{m_R(v)_{t+1}} g$$

$$= y_{t+1} + f(\mathbb{1}_k \mathbb{V}^{1_R}) \mathbb{S}_{R_Y} \mathbb{S}_{K_R} (R_{t+1}^{1_R})$$

(77)

and for $\pm^2$:

$$\pm^2 = f(\mathbb{1}_k \mathbb{V}^{1_R}) \mathbb{S}_{R_Y} \mathbb{S}_{K_R} f(\mathbb{1}_k \mathbb{V}^{1_R}) \mathbb{S}_{R_Y}$$

(78)
For further reference it is useful to define the vector \( v \) as

\[
(v) = (v) \quad \text{and} \quad (v) = \frac{1}{2} f(\phi_k i \ v^1_i) \quad \text{and} \quad (v) = R_y g. \tag{79}
\]

Notice that the expressions for \( (v) \) and \( \hat{\mu} \) given here differ slightly from the ones given in Hansen and Jagannathan (1997) because we explicitly included a constant in the projections of \( M (v)_{t+1} \) and \( y_{t+1} \) on \( R_{t+1} \).

The expressions for \( m_R(v)_{t+1} \) and \( \hat{\mu} \) in (77) and (78) provide a basis to relate the specification error bounds to intersection. In case of intersection the interest is in stochastic discount factors that price both \( R_{t+1} \) and \( r_{t+1} \), i.e., in \( M (v)_{t+1} \). Therefore, in the expressions (77) and (78) we should leave out all the \( R \)-subscripts, replace \( R_{t+1} \) with the vector \((R^0_t r^0_t)^0\), and note that all vectors and matrices have dimension \( K + N \) rather than \( K \). As before, with intersection we want to know if the minimum variance stochastic discount factor based on \( R_{t+1} \) only, \( m_R(v)_{t+1} \) can be used to price both \( R_{t+1} \) and \( r_{t+1} \). In terms of specification errors this means that we want to use \( m_R(v)_{t+1} \) as a proxy \( y_{t+1} \) for the stochastic discount factors \( M (v)_{t+1} \). Also, in the spirit of the previous section, when using \( m_R(v)_{t+1} \) as a proxy, we recognize beforehand that \( m_R(v)_{t+1} \) will not assign the correct prices to \( r_{t+1} \), but the interest is in the extent to which the assigned prices are wrong, i.e., the extent to which there are deviations from intersection, as measured by \( \pm \).

Recall that the proxy \( y_{t+1} = m_R(v)_{t+1} \) is now given by

\[
y_{t+1} = m_R(v)_{t+1} = v + \left( R(v) \right)^q R_{t+1} ^1 R; \quad \left( R(v) \right)^q R_{t+1} ^1 R \tag{80}
\]

Substituting these expressions into (77) and (78), properly adjusted for the fact that the interest is now in stochastic discount factors that price both \( R_{t+1} \) and \( r_{t+1} \), straightforward algebra shows that

\[
\hat{\mu} = f(\phi_k i v^1_i) \quad \text{and} \quad \hat{\mu} = \frac{1}{2} f(\phi_k i v^1_i) g^2 \frac{1}{2} f(\phi_k i v^1_i) \quad \text{and} \quad \hat{\mu} = \frac{1}{2} f(\phi_k i v^1_i) g^2 \frac{1}{2} f(\phi_k i v^1_i) \tag{80}
\]

or

\[
\hat{\mu} = f\mu(1=v)^2 \quad \text{and} \quad \hat{\mu} = (1=v)^2 g^2 \frac{1}{2} f(\phi_k i v^1_i) \tag{80}
\]

where \( \hat{\mu} \) is the covariance matrix of the residuals \( y_{t+1} \) from a regression of \( r_{t+1} \) on \( R_{t+1} \) and a constant. Also, the stochastic discount factor closest to \( y_{t+1} \) is now given by

\[
m_R(v)_{t+1} = \frac{1}{2} f(\phi_k i v^1_i) \quad \text{and} \quad \hat{\mu} = (1=v) \frac{1}{2} f(\phi_k i v^1_i) \tag{80}
\]

or

\[
m_R(v)_{t+1} = \frac{1}{2} f(\phi_k i v^1_i) \quad \text{and} \quad \hat{\mu} = (1=v) \frac{1}{2} f(\phi_k i v^1_i) \tag{80}
\]

where \( \hat{\mu} \) is the covariance matrix of the residuals \( y_{t+1} \) from a regression of \( r_{t+1} \) on \( R_{t+1} \) and a constant. Also, the stochastic discount factor closest to \( y_{t+1} \) is now given by

\[
m_R(v)_{t+1} = \frac{1}{2} f(\phi_k i v^1_i) \quad \text{and} \quad \hat{\mu} = (1=v) \frac{1}{2} f(\phi_k i v^1_i) \tag{80}
\]

or

\[
m_R(v)_{t+1} = \frac{1}{2} f(\phi_k i v^1_i) \quad \text{and} \quad \hat{\mu} = (1=v) \frac{1}{2} f(\phi_k i v^1_i) \tag{80}
\]
Thus, if we want to use the stochastic discount factor that is on the volatility bound of $R_{t+1}$, as a proxy stochastic discount factor for the larger set $(R_{t+1}; r_{t+1})$, then the valid discount factor that is closest to $m_R(v)_{t+1}$ is the discount factor with the same expectation $v$ that is on the volatility bound of $(R_{t+1}; r_{t+1})$. Therefore, $\pm$ is the least squares distance between two stochastic discount factors that are on the volatility bounds of $(R_{t+1}; r_{t+1})$ and its sub-set $R_{t+1}$ respectively, and is a straightforward measure of the deviation from intersection, which shows the close relation between this special case of the specification error bound and intersection. This relationship also follows from (80), which shows that $\pm$ is directly related to the change in the maximum squared Sharpe ratios that can be attained with $R_{t+1}$ and $(R_{t+1}; r_{t+1})$ respectively. It also follows that $\pm$ measures the difference between the variances of the two minimum variance kernels: $\pm = \text{Var}(m_R(v)_{t+1}) - \text{Var}(m_R(v)_{t+1})$.

An estimate of $\pm^2$ can easily be obtained from the sample equivalent of (78), which we will denote by $\hat{\pm}^2$. If the interest is in whether or not there is intersection, then we want to know whether or not $\pm = 0$, and this hypothesis can easily be tested as outlined in Section 3. From the expression in (80) and the discussion in previous sections, it follows that under the null hypothesis that $\pm = 0$,

$$T \frac{\hat{\pm}^2}{\sigma^2(1 + \beta R(1-v)^2)} \sim \chi^2_N.$$  \hfill (82)

In case of specification errors however, the interest is in the case where $\pm$ is strictly positive rather than zero. For that case the limiting distribution of $\hat{\pm}$ is derived in Hansen, Heaton, and Luttmer (1995).

Once we concede that $y_{t+1} = m_R(v)_{t+1}$ is not a valid stochastic discount factor for $(R_{t+1}; r_{t+1})$, we want to have a measure of the difference between $m_R(v)_{t+1}$ and the valid stochastic discount factor that is closest to it, $m_R(v)_{t+1}$. The specification error bound $\pm$ is one such measure, allowing us to make statements about how good or how bad the proxy performs. The fact that $\pm^2$ is equal to the change in maximum Sharpe ratios, makes the measure $\pm$ also useful in terms of the optimal portfolio choice for a mean-variance investor. Recall that a mean-variance investor that initially only invests in $R_{t+1}$ can improve his Sharpe ratio from $\mu_R(1-v)$ to $\mu(1-v)$ by including $r_{t+1}$ in his portfolio. Given that there is no intersection between the mean-variance frontiers of $R_{t+1}$ and $(R_{t+1}; r_{t+1})$, $\hat{\pm}$ provides an estimate for the potential increase in Sharpe ratios. Notice though that such an estimate can also be derived directly from the Wald test-statistic for intersection.
7 Summary

The purpose of this paper is to analyze and illustrate the concept of mean-variance spanning and intersection. We show that there is a duality between mean-variance frontiers and volatility bounds and that mean-variance spanning and intersection can be understood both in terms of mean-variance frontiers and volatility bounds. The paper shows how regression based tests can be used to test for spanning and intersection and how these regression based tests are related to tests for mean-variance efficiency, performance measurement, optimal portfolio choice and specification error bounds.

A The graphical relationship between mean-variance frontiers and volatility bounds

In this appendix we will show some graphical relations between the volatility bound and the mean-variance frontier for a set of asset returns $R_{t+1}$ with expectation $¹$ and covariance matrix $§$. We will start from a point on the volatility bound where the expectation of the minimum variance pricing kernel is $v$, i.e.,

$$E[m(v)_{t+1}] = v;$$

(83)

Using the efficient set variables $A$, $B$, and $C$, and the variance of $m(v)_{t+1}$ as given in (7), the variance of $m(v)_{t+1}$ can be written as

$$\text{Var}[m(v)_{t+1}] = A \cdot 2B \cdot v + C \cdot v^2;$$

(84)

which is a simple quadratic function of $v$ that describes the volatility bound. The second panel of Figure 1 gives a plot of $\text{Var}[m(v)_{t+1}]$ as a function of $v$.

As shown in Section 2.2, each minimum variance pricing kernel $m(v)_{t+1}$ corresponds to a mean-variance efficient portfolio that has a zero-beta rate $'= 1=v$. Recall that a mean-variance efficient portfolio satisfies

$$w = \delta_i 1_S i \lambda^{(1)} i \cdot \emptyset;$$

for a given risk aversion $\delta$ and associated zero-beta rate $'$. Using $\emptyset w = 1$ it follos that

$$\delta = B i \cdot A:$$
Furthermore, the expected portfolio return $^{1}w$ satisfies

$$^{1}w = ^{0}w \left( ^{C}i \cdot ^{B} \right) = \frac{^{C}i \cdot ^{B}}{^{B}i \cdot ^{A}};$$

Denote the return on the mean-variance efficient portfolio with zero-beta rate $^{1}w$ as $R(v)_{t+1}$ and define $^{1}w = E[R(v)_{t+1}]$. From the previous relations $^{1}w$ can be written as a function of $v$:

$$^{1}w(v) = \frac{^{B}i \cdot ^{C}v}{^{A}i \cdot ^{B}v};$$  \hspace{1cm} (85)

Also, the variance $w^{2}$ for a mean-variance efficient portfolio $w$ can be written as a function of $^{1}w(v)$:

$$\text{Var}[R(v)_{t+1}] = \frac{^{A}i \cdot ^{2}B \cdot ^{1}w(v) + C}{^{A}i \cdot ^{B}v};$$

or as a function of $v$:

$$\text{Var}[R(v)_{t+1}] = \frac{^{A}i \cdot 2^{B}v + ^{C}v^{2}}{^{A}i \cdot ^{B}v^{2}};$$  \hspace{1cm} (86)

The first panel of Figure 1 shows the standard mean-variance efficient frontier, where the expected portfolio return $^{1}w(v)$ is plotted as a function of the standard deviation of the portfolio return $\text{std}(R(v)_{t+1}) = \sqrt{\text{Var}[R(v)_{t+1}]}$. In this appendix we will restrict ourselves to characterizing the relation between the volatility bound and the mean-variance frontier in terms of $v$ and $^{1}w(v)$. Given the relations (84) to (86) above it is straightforward to derive the variances of the pricing kernel and the associated mean-variance efficient portfolio as well.

To see the relation between the two graphs, all notice that the expected portfolio return $^{1}w(v)$ is decreasing in $v$, since from (85) we have that

$$\frac{\partial (^{1}w)}{\partial v} = \frac{^{B}i \cdot ^{A}C}{(^{A}i \cdot ^{B}v)^{2}} < 0;$$

and where the inequality follows from the fact that $^{A}C > ^{B}^{2}$, by the Cauchy-Schwarz inequality (see also Ingersoll (1987, p.85)).

Next, from (85) it also follows that for $v = 0$ we have that $^{1}w = ^{B} \Rightarrow ^{A}$, which is the expected return on the Global Minimum Variance portfolio.
Looking at the volatility of the pricing kernel we can of course also distinguish the Global Minimum Variance Pricing Kernel, the expectation of which can be found using (84):

\[ 0 = \frac{\text{Var}[m(v)_{t+1}]}{\text{Var}} = i \cdot 2B + 2C \nu^* \]

The second derivative \(2C\) is always positive, which confirms that this is indeed a minimum. Using (85) again, \( v = B/C \) corresponds to \( ^1(\nu) = 0 \). Thus, when the expectation of the kernel is zero; \( v = 0 \), this corresponds to the Global Minimum Variance portfolio on the mean-variance frontier, whereas a zero expected return for the mean-variance efficient portfolio, \( ^1(\nu) = 0 \), in turn corresponds to the Global Minimum Variance kernel on the volatility bound.

Having characterized the global minima of the two frontiers, the next step is to look at the other extremes, i.e., where \( v \to -1 \) and where \( ^1(\nu) \to -1 \).

Taking limits and using (85) we get that

\[ \lim_{v \to -1} \frac{B}{A} \frac{Cv}{Bv} = \frac{C}{B} \]

Thus, both extremes of the left and right limb of the volatility bound correspond to the same single point on the mean-variance frontier, where the expected portfolio return is \( ^1(\nu) = C/B \). Since by the Cauchy-Schwarz inequality \( C/B > B/A \) if \( B > 0 \), the point where \( ^1(\nu) = C/B \) will plot on the upper limb of the mean-variance frontier. \( B > 0 \) is the typical case, since this implies that with positive interest rates or zero-beta returns, efficient portfolios have positive expected returns. It is useful to note that \( ^1(\nu) = C/B \) corresponds to the point where a straight line through the origin is tangent to the mean-variance frontier (since \( v \to -1 \) corresponds to \( ^1(\nu) \to 0 \)).

Finally, by rewriting (85) as

\[ v = \frac{B}{C} \frac{A^1(\nu)}{B^1(\nu)} \]

we can find the point(s) on the volatility bound that correspond to the extremes of the mean-variance frontier, i.e., where \( ^1(\nu) \to -1 \). Taking limits
again, we get that

\[
\lim_{v \to 1} \frac{B}{C} \frac{A^1(v)}{B^1(v)} = \frac{A}{B} ; \\
\lim_{v \to 1 +} \frac{B}{C} \frac{A^1(v)}{B^1(v)} = \frac{A}{B} .
\]

Notice that we already discussed this result in Section 2 since \( v = A = B , \) \( \acute{v} = B = A , \) i.e. the case where the zero-beta return equals the expected return on the Global Minimum Variance portfolio and where there are no corresponding mean-variance efficient portfolios, since the asymptotes of the mean-variance frontier cross the y-axis at \( B = A , \) but there is no line tangent to the frontier starting at this point. Again, if \( B > 0 , \) then the Cauchy-Schwarz inequality implies that \( A = B > B = C , \) implying that this point will be located on the right limb of the volatility bound. Finally, it is useful to note that if we would plot the volatility bound as the standard deviation of the pricing kernel, \( \text{Var}[m(v)_{t+1}] , \) as a function of \( v , \) then \( v = A = B \) would correspond to the point where a straight line through the origin is tangent to the volatility bound, similar to the mean-variance frontier when \( v = C = B . \)

B. The spanning test-statistic in terms of Sharpe ratios

In this appendix we show how the spanning test statistic can be interpreted in terms of Sharpe ratios, a result that was presented in Section 5.3. Recall from Section 5.3 that the covariance matrix of the OLS-estimates \( \hat{B} \) equals

\[
\hat{A} = \lim_{t \to 1} 1 + \sum_{i} \frac{1}{R} \frac{i R}{R^1} \frac{i R}{R^1} \frac{1}{R} .
\]

Premultiplying with \( H_{\text{span}} \) and postmultiplying with \( H_{\text{span}} \) as defined in (55) yields

\[
H_{\text{span}} \lim_{t \to 1} 1 + \sum_{i} \frac{1}{R} \frac{i R}{R^1} \frac{i R}{R^1} \frac{1}{R} = \lim_{t \to 1} 1 + C + \sum_{i} \frac{1}{R} \frac{i R}{R^1} \frac{i R}{R^1} \frac{1}{R} .
\]

(87)
the inverse of which is
\[ \frac{x_i,1}{A_R(1 + C_R)} i \]

\[ A_R B_R \equiv 1 : \] (88)

Similarly, for \( h_{\text{span}} \) in (55) we have
\[ i N - 1 \quad 0 \quad b_i \quad i N - 0 \quad \quad \begin{array}{c}
A \quad 0 \quad b_i \quad 1 \quad 0 \\
0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0 \\
\vdots \quad 0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0 \\
N \quad 0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0
\end{array} \]

\[ = 0 \quad \begin{array}{c}
0 \quad \begin{array}{c}
A \quad 0 \quad b_i \quad 1 \quad 0 \\
0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0 \\
\vdots \quad 0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0 \\
N \quad 0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0
\end{array}
\end{array} \]

Premultiplying (88) with \( h_{\text{span}} \) and postmultiplying with \( h_{\text{span}}^0 \), we get, after replacing population moments by their sample equivalents:
\[ \begin{pmatrix} 0 \quad 1 \quad 0 \quad b_i \quad 0 \quad 0 \quad b_i \quad 1 \quad 0 \\
A \quad 0 \quad b_i \quad 1 \quad 0 \\
\vdots \quad 0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0 \\
N \quad 0 \quad b_i \quad 0 \quad b_i \quad 1 \quad 0
\end{pmatrix} \]

(89)

Next note that the maximum attainable Sharpe ratio from \( R_{t+1} \), for \( B_R = A_R \), is equal to
\[ \mu^\prime \frac{B_R}{A_R} \]

For simplicity, write \( A = A_R + \xi A \), \( B = B_R + \xi B \), and \( C = C_R + \xi C \), where the definitions of \( \xi A \), \( \xi B \), and \( \xi C \) follow from (54) and (55). Evaluating \( \mu^\prime \) in this same value of \( \xi \), we get
\[ \mu \frac{B_R}{A_R} \frac{\xi_2}{A_R} = C_i \frac{B_R}{A_R} + A_R \frac{B_R^2}{A_R} \]

\[ = C_R + \xi C_i \frac{B_R}{A_R} + A_R \frac{B_R^2}{A_R} \]

\[ = \mu_R \frac{B_R}{A_R} \frac{\xi_2}{A_R} + \frac{1}{A_R} A_R \xi C_i \frac{B_R}{A_R} + \frac{B_R^2}{A_R} \xi A \]

Dividing by \( (1 + C_R) i \), \( B_R = A_R \), \( 1 + \mu R \frac{B_R}{A_R} \frac{\xi_2}{A_R} \) gives
\[ \frac{\mu^3 \frac{B_R}{A_R} \frac{\xi_2}{A_R} \frac{B_R}{A_R} \frac{\xi_2}{A_R}}{1 + \mu_R \frac{B_R}{A_R} \frac{\xi_2}{A_R}} = \frac{A_R \xi C_i \frac{B_R}{A_R} + \frac{B_R^2}{A_R} \xi A}{A_R (1 + C_R) i \frac{B_R^2}{A_R}} \]

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Replacing all population moments with their sample equivalents again and noting that \( 1 = \lambda_R \) is the variance of the global minimum variance portfolio of \( R_{t+1} \), i.e., \( 1 = \lambda_R = (\hat{\beta} R)^2 \), and similarly, \( 1 = \lambda = (\hat{\beta} R)^2 \), we finally obtain

\[
\begin{align*}
\text{span}_W & = T \frac{\beta^3 \frac{\lambda R^2}{\lambda R} \beta}{1 + \beta R} i \frac{\beta^3 \frac{\lambda R^2}{\lambda R} \beta}{1 + \beta R} + T \lambda R \frac{\lambda R}{\lambda R^2} i \\
& = T \frac{1 + \beta (\hat{\beta} R)^2}{1 + \beta \hat{R} (\hat{\beta} R)^2 + (\hat{\beta} R)^2} \lambda^2 i 2 \end{align*}
\]
References


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91.


