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Systematic Risk, Hedging Pressure, and Risk Premiums in Futures Markets

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I examine the uniformity of risk pricing in futures and asset markets. Tests against a general alternative do not reject complete integration of futures and asset markets. As predicted, estimates of the "zero-beta" rate for futures are close to zero, and premiums for systematic risk do not differ significantly across assets and futures. There is, however, evidence consistent with a specific alternative model presented by Hirshleifer (1988). Returns in foreign currency and agricultural futures vary with the net holdings of hedgers, after controlling for systematic risk. These results imply a degree of market segmentation and support hedging pressure as a determinant of futures premiums.

Futures markets facilitate the transfer of risk to those most able or willing to bear it. The market price, or risk premium, for this transfer is the difference between the current futures price and the expected price on the same contract at a later date. Numerous empirical studies document that futures prices do not follow random walks.[1] However, studies such as those

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[1] Examples include Bodie and Rosansky (1980) and Fama and French (1987) for commodity futures, Hodrick and Srivastava (1987) and Bilson (1981) for foreign currency futures, and Rzepczynski (1987) and Goldenberg (1988) for financial futures.

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by Jagannathan (1985) and Hodrick and Srivastava (1983) that test equilibrium models of futures risk premiums generally reject the model in favor of an unspecified alternative, leaving no empirically supported model of futures premiums.

In this study, I use a comprehensive sample drawn from financial, foreign currency, metal, and agricultural futures markets to test a simple model of futures risk premiums. The model, based on complete integration of asset and futures markets, is tested against a general alternative by examining the implications that (1) the "zero-beta" rate for futures is equal to zero and (2) premiums for systematic risk in futures and assets are equal. I also test the model against a specific alternative provided by Hirshleifer (1988), in which the futures pricing function contains hedging-dependent (and possibly market-specific) premiums for residual risk in addition to premiums for systematic risk.

Results of tests conducted against the general alternative do not reject the hypothesis of full integration. Estimates of the zero-beta rate for futures are close to zero, and premiums for systematic risk in futures markets do not differ significantly from premiums in asset markets. However, for some futures markets the hypothesis of full integration is rejected in favor of Hirshleifer's alternative equilibrium model. Specifically, mean returns in foreign currency and agricultural futures differ significantly, as a function of the net holdings of hedgers, from the level predicted based on systematic risk alone. Thus, the results indicate a degree of segmentation between asset markets and some futures markets, and support net hedging pressure as a determinant of risk premiums in those markets.

1. Determinants of Futures Risk Premiums

Theoretical analysis identifies two determinants of futures premiums: hedging pressure and systematic risk. A class of models dating to Keynes (1930) and Hicks (1939) focuses on traders who participate in futures markets to reduce risk. Their net supply of futures contracts, or "hedging pressure," causes the equilibrium path of futures prices to contain an expected time trend. In Keynes' (1930) original theory, hedgers are endowed with long positions in the underlying good and thus take short positions in futures, resulting in an equilibrium where futures prices tend to rise over time. More recent work [e.g., Cootner (1960)] recognizes that hedgers can be purchasers as well as producers of the underlying good, and can face quantity as well as price risk [e.g., McKinnon (1967), Rolfo (1980), and Anderson and Danthine (1983)]. In these models, hedgers may take long positions in futures markets, resulting in an equilibrium where futures prices tend to decrease over time. These models imply that the sign of the futures risk premium varies cross-sectionally and intertemporally as a function of the net holdings of hedgers.

The literature on hedging pressure typically assumes an economy in which a small number of securities are traded. A key implicit assumption is that hedgers cannot market claims to their profits. In contrast, models based on the assumption that all claims are costlessly marketable (so portfolios can be freely diversified) imply that futures risk premiums are functions of nondiversifiable risk rather than hedging pressure. In these models, futures premiums depend on the degree of covariation between futures prices and changes in economic state variables, without a role for hedging pressure. [See Dusak (1973), Black (1976), Richard and Sundaresan (1981), and Hodrick and Srivastava (1987).]

There is, however, empirical evidence that movements in futures prices are correlated with hedgers' positions. Chang (1985) examines whether price movements in three agricultural futures markets are related to the net positions of large hedgers and large speculators, as reported in Commodity Futures Trading Commission (CFTC) publications. He finds that prices rise more often than expected on a random basis in months when large speculators have net long positions and fall more often than expected in months when large hedgers have net long positions. Carter, Rausser, and Schmitz (1983) report that, for five agricultural futures, intercepts in regressions of futures returns on a proxy for the market portfolio vary as a function of CFTC reported net hedging.

Hirshleifer (1988, 1990) presents equilibrium models that integrate modern portfolio theory and the insights of earlier theoreticians. His models incorporate two market imperfections. First, some claims are nonmarketable, as in Mayers (1972). Second, in the spirit of Merton's (1987) model of the effects on nonparticipation in security markets, participation in futures markets is limited by the existence of fixed (possibly informational) setup costs. In the resulting equilibrium, futures risk premiums depend on both systematic risk and residual risk, and the sign of the residual risk premium depends on the sign of net hedging pressure.

In this study, I provide empirical evidence on the roles of systematic risk and hedging pressure in determining futures risk premiums. Unlike prior studies, which typically focus on a closely related group of futures, this study uses prices drawn from a cross section of 22 financial, foreign currency, agricultural, and metal futures markets from 1967 (or the origination of trading) to 1989, in addition to data on equity returns. A broad cross section is desirable for two reasons. First, as results in Stambaugh (1982) demonstrate, inference with regard to asset pricing models can be sensitive to the exclusion of securities from the cross-sectional analysis. Second, since the importance of market imperfections, such as those incorporated by Hirshleifer (1988), may well vary across types of futures, a broad cross section allows examination of whether risk pricing varies across futures markets.

2. The Implications of Market Integration

Typical definitions of capital market integration imply that financial assets that trade in different markets but that have identical risk characteristics will have identical expected returns [e.g., Campbell and Hamao (1992)]. For futures markets, this definition must be modified to reflect that, unlike asset markets, the market value of a futures position is zero at origination.[2] Integration of asset and futures markets is taken here to imply that expected returns on portfolios consisting of asset and futures positions are identical to expected returns on asset-only portfolios of identical risk.

Conditional versions of modern asset pricing theories imply a linear relation between expected return and systematic risk, stated as

$$E_{t-1}(R^a_{it}) = \gamma_{0t} + \beta^a_{it-1}\gamma_{1t}, \tag{1}$$

where $E_{t-1}(R^a_{it})$ is expected next period return on asset i at time $t-1$, γ_{0t} is a cross-sectional constant (the zero-beta rate at time t), β^a_{it-1} is a $1 \times p$ vector of conditional sensitivities (multiple regression coefficients) of asset i returns to each of p economic aggregates, and γ_{1t} is a $p \times 1$ vector of risk premiums in effect at time t.[3]

It is easily shown that market integration implies that for zero-investment positions, such as futures, (1) is modified to exclude the intercept. That is, the expected percentage change in the futures price, $E_{t-1}(R^f_{jt})$, obeys the relation

$$E_{t-1}(R^f_{jt}) = \beta^f_{jt-1}\gamma_{1t}, \tag{2}$$

where β^f_{jt-1} is a $1 \times p$ vector of conditional sensitivities of percentage

[2] Because of the zero capital investment, rates of return on futures positions are not defined in the usual sense. Futures traders must typically provide capital for initial margin, but margin merely serves to bond performance under the contract, and is not a purchase price. Further, margin can generally be provided in the form of interest-bearing instruments, with interest accruing to the futures trader.

[3] A conditional version of the capital asset pricing model of Sharpe (1964) and Black (1972) implies a single-beta version of Equation (1), with beta measuring covariation of returns with returns to the market portfolio. Multiple-beta versions of (1) can be motivated by use of a multivariate proxy for the market portfolio, or by use of proxies for the multiple state variables in models such as Merton (1973), Ross (1976), and Connor (1984). In each of these models, the fundamental economic aggregate is the marginal utility of the representative investor.

changes in futures price j to the p economic aggregates, and γ_{1t} is defined as above.[4]

Equation (2) defines a conditional pricing function for futures that is parallel to the conditional pricing function for assets defined by (1), but (2) has an intercept equal to zero. The equality of risk premiums across (1) and (2) implies a uniform trade-off of portfolio risk for expected return across asset and futures markets, while the zero intercept reflects the zero capital investment in a futures position. If Equation (1) holds for assets and Equation (2) holds for futures, then expected returns on portfolios of assets and futures are also given by (1). Also, given (1), rejection of (2) implies the existence of asset-futures portfolios that dominate asset-only portfolios in terms of expected return and beta risk, which is characteristic of market segmentation. This analysis also formalizes the relation between the risk premium in a futures market and the expected excess return (over the zero-beta rate) in the associated spot market. Market integration implies that the futures premium and the expected excess return on the spot asset differ only if the systematic risk of the spot and the future differ.

Equation (2) relies on the integration of asset and futures markets as well as on the validity of the asset pricing relation (1). Tests of (2) are joint tests of integration and of the other assumptions of the underlying asset pricing theory. Rejection of (2) could be due to segmentation of markets, failure of the asset pricing theory, or improper measurement of systematic risk.

3. Data

Daily settlement prices for 22 futures markets, from January 1967 (or the origination of contract trading) to December 1989, are obtained from the Center for the Study of Futures Markets at Columbia University and from Data Resources, Inc. The data include six financial futures (Treasury bonds, Treasury notes, Treasury bills, Eurodollars, S&P 500 Index, and Value Line Composite Index), six agricultural futures (live cattle, soybeans, sugar, wheat, cotton, and corn), five foreign currency futures (Canadian dollar, British pound, Japanese yen, German mark, and Swiss franc), and five mineral futures (gold, silver, platinum, copper, and crude oil). These markets are selected for the cross section of different underlying goods they provide, and based on their economic importance as evidenced by relatively large trading volumes.

[4] The similarity between Equations (1) and (2) illustrates the analogy between rates of change in futures prices and rates of return on assets. Though a misnomer, percentage changes in futures prices are hereafter referred to as futures returns.

Individual futures contracts have a finite life defined by the contractual delivery date. Futures return series are compiled as daily percentage changes in the settlement price of the contract with the nearest delivery date, except within the delivery month when daily percentage changes in the settlement price of the second-nearest contract are used.[5] Monthly futures returns are obtained by linking these daily returns.

Price limits introduce a source of error in measuring futures returns. The incomplete adjustment of reported prices in the presence of a limit move introduces effects similar to those of nonsynchronous trading on measured portfolio returns. In particular, bias can be introduced into estimates of systematic risk. While it is difficult to assess the sensitivity of reported results to this source of measurement error, two considerations suggest the effects are minimal. First, limit moves are relatively rare. For example, data in Kodres (1988) indicate that limit moves in foreign currency futures affect less than 1 percent of observations. Second, systematic risk is estimated on the basis of monthly rather than daily returns. This reduces the relative magnitude of errors and, since no error is introduced unless a limit move occurs on the last day of the month, it reduces the likelihood that the returns used for beta estimation contain errors.

I examine the pricing of systematic risk using both single-beta and multiple-beta models. Single-beta models are estimated using returns to a diversified equity portfolio, the CRSP value-weighted index, as the benchmark. Multiple-beta models are estimated using, in addition to the CRSP index returns, six macroeconomic variables similar to those used by Chen, Roll, and Ross (1986): unexpected inflation, change in expected inflation, change in short-term Treasury-bill yields, change in the term structure (long-term government bond yield less Treasury-bill yield), change in default premium (yield on bonds rated BAA by Moody's less the long-term government bond yield), and unexpected change in U.S. industrial production. Data sources and construction of these variables are described in the Appendix.

Data on net hedging are obtained from various issues of *Commitments of Traders in Commodity Futures,* published by the CFTC. All large traders are required to identify themselves as hedgers or speculators. The CFTC publication reports, by market, on a monthly basis,

[5] Each return is computed using successive prices on a contract for delivery on a specific date, and never across contracts with different delivery dates. Returns reflect a strategy of closing the position in the near contract and opening a position in the second-nearest contract at the end of trading the month before delivery. The quoted settlement price in some markets does not reflect the delivery price. For example, a quoted Treasury-bill price of QP indicates that sellers are obligated to deliver (at contract maturity) 90-day Treasury bills for a price such that the annualized return to the buyer is 100 − QP percent. Quoted prices are converted to implied delivery prices, and returns computed as the percentage change in these implied delivery prices.

the long open interest and short open interest held by traders reported to be hedgers. If reported hedgers' short open interest exceeds reported hedgers' long open interest, I identify the market as being characterized by net short hedging and long speculation, and vice versa.[6]

Hedgers' open interest is reported as of month-end. Prior researchers, including Chang (1985) and Carter, Rausser, and Schmitz (1983), take month-end hedging as indicative of positions throughout the month. This is problematic when the sign of net hedging differs from beginning to end of month. To be more conservative, I identify hedging pressure as short or long only in months where reported net hedging does not change sign, and evaluate separately months during which net hedging changes sign. For three metals markets (silver, platinum, and copper) long hedging exceeds short hedging, implying net long speculation, in all months for which data are available. For a nontrivial number of observations, data on net hedging are not obtainable, either because the CFTC reports were not published or because data were omitted from the published reports.[7] Months for which hedging data are not available are also evaluated separately.

4. Mean Returns and Hedging Pressure in Futures

Mean daily returns (annualized by multiplying by 250) are displayed for each of the 22 futures markets and five portfolios of futures positions in Table 1. Means are reported for the full sample and for subsamples partitioned on the basis of hedgers' positions. Portfolio means are averages of the time series of daily cross-sectional portfolio means, where daily means are computed using returns in those futures markets open for trading on that day.[8]

Except for some financial futures, there is little evidence that unconditional mean futures returns differ from zero. The hypothesis that

[6] This convention implicitly categorizes small (nonreporting) positions as speculative. While older CFTC reports explicitly label traders as hedgers or speculators, more recent reports use the categories "commercial" and "noncommercial." Noncommercial traders are likely to be speculators in the theoretical sense. Commercial traders plausibly include hedgers, but some commercial traders may take speculative positions. Therefore, the identification of commercial positions as hedging positions is likely to contain some noise.

[7] There are two main types of omissions. First, hedging data are typically not included in CFTC publications until futures markets are opened for a period of time. Thus, for most markets hedging data are missing for earlier years of the sample. Second, the CFTC did not publish hedging data from January to November 1982. Information on the number of missing observations by market is provided in Table 1. In aggregate, contemporaneous hedging data are available for 68 percent of the return observations in the sample.

[8] The number of markets open for trading, and hence the number of returns used to compute daily portfolio means, increases over time. This may introduce heteroskedasticity into the time series of daily portfolio means. To assess sensitivity to heteroskedasticity, I also compute time-series means by weighted least squares regressions of daily portfolio means on a constant, using the number of markets open each day as the weighting variable. The resulting t-statistics are little altered, and conclusions are unaffected.

the unconditional mean return is zero is rejected in only one (live cattle) of the 16 nonfinancial futures markets. Unconditional mean returns on equal-weighted portfolios of futures are small in economic terms as well: 2.4 percent in agricultural commodities, 4.5 percent in metals, and 0.7 percent in foreign currencies. However, in financial futures unconditional mean returns on Treasury notes and Eurodollars are statistically significant, and unconditional mean returns on S&P 500 Index futures (12.8 percent) and Value Line Index futures (9.1 percent) are economically, though not statistically, significant.

In contrast to results for unconditional means, substantial evidence indicates that mean returns on nonfinancial futures differ from zero when conditioned on the sign of net hedging. Conditional on short hedging (long speculation), mean returns are significantly positive in seven of the 16 nonfinancial futures markets. Conditional on long hedging (short speculation), mean returns are significantly negative in seven of the 13 nonfinancial markets for which a conditional mean can be computed. The right column of Table 1 shows a test statistic for the hypothesis that mean returns conditional on net long speculation are equal to mean returns conditional on net short speculation.[9] This hypothesis is rejected ($p < .01$) for eight of the 13 nonfinancial markets in which the test can be conducted. In each case where the hypothesis is rejected, the mean return conditional on long speculation exceeds the mean return conditional on short speculation.

Multivariate evidence is obtained by examining time-series means of daily returns to equal-weighted portfolios. Each daily hedging-conditioned portfolio return is computed using those markets open for trading and characterized by the appropriate hedging status. Conditional on short hedging (long speculation), the mean return to each of the three nonfinancial portfolios is positive. Conditional on long hedging (short speculation), the mean return on each of these three portfolios is negative. The hypothesis that the mean return conditioned on long speculation equals the mean return conditioned on short speculation is rejected for each of the agricultural, foreign currency, and metal portfolios, as well as for an equal-weighted portfolio of all 22 futures markets.

These findings are broadly consistent with traditional net hedging pressure theories and the prior empirical evidence on hedging pressure. Chang (1985) and Carter, Rausser, and Schmitz (1983) report results only for agricultural commodities. The evidence here indicates

[9] I regress futures returns on appropriately defined dummy variables and test whether coefficient estimates on the dummies are equal. The test uses the White (1980) heteroskedasticity consistent covariance matrix, resulting in a test statistic distributed χ^2.

that hedging pressure also has explanatory power in other futures markets. The evidence is strongest in foreign currency futures, where mean returns conditional on net long speculation exceed mean returns conditional on net short speculation by an economically and statistically significant margin for all five foreign currencies.

In contrast to results for nonfinancial futures, the effect of hedging pressure on financial futures is mixed, and often opposite of that predicted by hedging pressure theories. For Treasury bills and Eurodollars, mean returns conditional on short hedging exceed mean returns conditional on long hedging. For the other four financial futures, and for an equal-weighted portfolio of financial futures, the opposite result holds. These results suggest that the market imperfections implicit in hedging pressure theories are not relevant to financial futures markets.

5. Systematic Risk in Futures Returns

The extent to which futures risk is systematic is crucial to predicting risk premiums in futures. Following Ferson and Harvey (1991), conditional betas for month t are estimated using data for the 60 months before t. As simple summary measures of the systematic risk in futures positions, estimates of unconditional betas for each market are reported in Table 2. Two sets of estimates are presented: "single beta," which are slope coefficients from simple regressions of monthly futures returns on monthly returns to the CRSP value-weighted equity index, and "multiple beta," which are slope coefficients from multiple regressions of monthly futures returns on returns to the CRSP value-weighted equity index and the six macroeconomic variables described above.

The beta estimates in Table 2 indicate that many futures positions are subject to nonzero systematic risk. As anticipated, returns on equity index contracts are positively related to spot equity returns, and returns on debt instrument contracts are negatively related to Treasury-bill yields and term structure variables. Returns to equity index futures are significantly related to macroeconomic variables even though the return on the CRSP equity portfolio is included in the multiple regression. Among foreign currency futures, only Canadian dollar returns are significantly related to U.S. equity returns, but the hypothesis that returns are independent of the six macroeconomic variables is rejected ($p < .10$) for each foreign currency. Returns to metal futures tend to be positively related to equity returns, although the relation is significant only for silver and platinum futures. Agricultural futures returns show the least evidence of systematic risk; only live cattle and soybean returns are significantly related to equity returns, and only

Table 1
Mean returns (% per day × 250) in futures (*t*-statistics in parentheses, sample sizes in brackets)

	Sample start date[1]	Unconditional	Conditional on net hedging[2]				Test: Long = short χ^2 statistic (*p*-value)
			Short	Long	Change	Missing	
Financial futures							
Treasury bills	1/76	0.12 (0.74) [3530]	0.09 (0.42) [2230]	−1.22 (−2.12)* [228]	−0.53 (−0.83) [234]	0.75 (2.57)** [838]	4.53 (.033)
Treasury bonds	9/77	1.99 (0.51) [3120]	−6.51 (−1.04) [1426]	1.60 (0.24) [843]	24.80 (2.53)* [424]	8.44 (0.92) [427]	0.77 (.379)
Treasury notes	5/82	7.54 (2.41)* [1938]	1.31 (0.31) [947]	20.42 (3.26)** [378]	2.08 (0.26) [296]	15.91 (1.58) [316]	6.51 (.011)
Eurodollars	5/82	0.53 (3.11)** [1938]	0.32 (1.44) [571]	0.15 (0.70) [928]	1.00 (2.58)** [270]	2.58 (2.19)* [169]	0.27 (.602)
S&P 500 Index	5/82	12.84 (1.61) [1938]	−2.26 (−0.14) [674]	25.38 (2.55)* [819]	3.81 (0.26) [298]	30.48 (0.87) [147]	2.03 (.154)
Value Line Index	5/82	9.09 (1.26) [1938]	7.02 (0.85) [1475]	37.80 (1.43) [108]	−9.65 (−0.58) [209]	35.41 (1.03) [147]	1.23 (.267)
Financial portfolio[3]		1.84 (0.86) [3530]	−2.22 (−0.66) [2465]	9.19 (2.49)* [1552]	3.80 (0.83) [1121]	6.91 (1.26) [838]	5.20 (.023)

Table 1
Continued

	Sample start date[1]	Unconditional	Conditional on net hedging[2] Short	Long	Change	Missing	Test: Long = short χ^2 statistic (*p*-value)
Agricultural futures							
Live cattle	1/67	7.88 (2.12)* [5792]	6.63 (1.36) [3883]	3.04 (0.05) [22]	32.85 (0.67) [21]	9.56 (1.79)* [1866]	0.01 (.954)
Soybeans	1/67	2.81 (0.57) [5788]	4.35 (0.66) [3015]	−1.21 (−0.11) [1506]	9.63 (0.89) [1057]	−24.84 (−1.42) [210]	0.18 (.669)
World sugar	1/67	−2.76 (−0.28) [5764]	−2.71 (−0.17) [2587]	71.88 (0.35) [21]	−11.82 (−0.76) [87]	−0.06 (−0.01) [3069]	0.13 (.717)
Wheat	1/67	0.92 (0.19) [5790]	5.71 (0.97) [4045]	−10.53 (−0.86) [569]	−5.93 (−0.53) [753]	−17.25 (−0.71) [423]	1.44 (.231)
Cotton	1/67	7.39 (1.64) [5746]	27.83 (4.18)** [2969]	−22.47 (−1.64) [645]	−22.86 (−1.22) [378]	−9.71 (−1.41) [1754]	10.91 (.001)
Corn	1/67	−1.56 (−0.38) [5805]	16.25 (2.69)** [2874]	−19.96 (−2.53)* [1630]	−15.79 (−1.76)* [1091]	−28.63 (−1.81)* [210]	13.25 (.000)
Agricultural portfolio[3]		2.39 (0.76) [5805]	10.93 (2.84)* [5290]	−9.62 (−1.47) [2894]	−7.86 (−1.32) [2412]	−8.47 −1.03 [3069]	7.29 (.009)

Table 1
Continued

	Sample start date[1]	Unconditional	Conditional on net hedging[2]				Test: Long = short χ^2 statistic (p-value)
			Short	Long	Change	Missing	
Metal and oil futures							
Gold	1/75	−0.40 (−0.06) [3773]	0.05 (0.00) [885]	−18.66 (−1.80)* [632]	−7.44 (−0.47) [269]	6.17 (0.62) [1987]	1.68 (.195)
Silver	1/67	2.91 (0.44) [5749]	−11.61 (−1.01) [1595]	— — [0]	— — [0]	8.38 (1.04) [4154]	—
Platinum	1/68	1.49 (0.23) [5514]	2.91 (0.31) [2678]	— — [0]	— — [0]	0.15 (0.02) [2836]	—
Copper	1/67	6.42 (1.15) [5754]	14.43 (1.44) [1808]	— — [0]	— — [0]	2.75 (0.41) [3946]	—
Crude oil	4/83	16.28 (1.28) [1697]	79.50 (2.67)** [567]	−13.67 (−1.17) [877]	−20.89 (−0.62) [253]	— — [0]	8.48 (.004)
Metal and oil portfolio[3]		4.53 (0.98) [5754]	5.31 (0.64) [2678]	−15.65 (−1.72)* [943]	−8.36 (0.58) [437]	5.91 (0.98) [4154]	2.71 (.088)

Table 1
Continued

	Sample start date[1]	Unconditional	Conditional on net hedging[2]				Test: Long = short χ^2 statistic (p-value)
			Short	Long	Change	Missing	
Foreign currency futures							
Canadian dollar	6/72	0.07 (0.08) [4443]	3.25 (2.03)* [1726]	−14.09 (−3.79)** [393]	2.55 (0.93) [570]	−0.69 (−0.56) [1754]	18.34 (.000)
British pound	6/72	0.20 (0.08) [4443]	14.72 (2.89)** [1334]	−22.81 (−3.25)** [693]	2.30 (0.31) [662]	−2.55 (−0.85) [1754]	18.77 (.000)
Japanese yen	6/72	1.76 (0.74) [4443]	9.63 (2.27)* [1684]	−27.80 (−3.15)** [379]	−6.31 (−0.85) [626]	3.49 (1.15) [1754]	14.65 (.000)
German mark	6/72	0.49 (0.20) [4443]	9.32 (1.66)* [1035]	−15.38 (−3.11)** [1136]	7.14 (0.77) [499]	3.62 (1.13) [1773]	10.86 (.001)
Swiss franc	6/72	1.01 (0.36) [4443]	3.11 (0.63) [1703]	−30.55 (−2.96)** [296]	−6.63 (−0.77) [670]	7.15 (1.90)* [1774]	8.64 (.003)
Foreign currency portfolio[3]		0.71 (0.40) [4443]	4.88 (1.75)* [2562]	−16.21 (−4.39)** [1681]	0.27 (0.07) [1558]	3.04 (1.41) [1774]	20.79 (.000)
All 22 futures portfolio[3]		2.71 (1.33) [5813]	9.68 (3.16)** [5291]	−6.03 (−1.32) [3941]	−2.02 (−0.52) [3188]	4.14 (0.97) [4237]	8.13 (.004)

[1] Each sample ends December 1989.
[2] Long denotes mean for months when hedgers are net long (speculators are net short) at beginning *and* end of month. Short denotes mean for months when hedgers are net short (speculators are net long) at beginning *and* end of month. Change denotes mean for months where the net position of hedgers differs at end of month from beginning of month. Missing denotes mean for months where hedging data are unavailable.
[3] Portfolio means reported are time-series averages of daily cross-sectional portfolio means, where each daily portfolio mean is computed using all markets open for trading on that date.
* Significantly different from zero with $p < .10$.
** Significantly different from zero with $p < .01$.

cotton returns are significantly related to the macroeconomic variables.

The estimates of systematic risk in Table 2 can be compared to those reported in earlier studies. Using data from 1952 to 1967, Dusak (1973) estimates equity index beta coefficients for wheat, corn, and soybeans that do not differ significantly from zero. Results obtained here for wheat and corn are similar to her estimates, but a significant negative coefficient for soybeans is found. Bodie and Rosansky (1980) report predominantly negative betas for agricultural and metal futures. Except for soybeans, the beta estimates obtained here are uniformly larger (near zero or positive) than the negative estimates they report. Breeden (1980) reports estimated consumption betas for agricultural and metal futures. Though measures of consumption are not included in this study, the results here, like Breeden's, indicate positive systematic risk for metals and generally zero systematic risk for grains.

As indicated above, the hypothesis tests in this study are based on conditional beta estimates rather than the unconditional estimates reported in Table 2. Possible time-series variation in the explanatory power of the economic aggregates is examined by comparing the time-series mean of the adjusted R^2 from successive 60-month rolling regressions (reported in square brackets in Table 2) with the adjusted R^2 from the unconditional regressions. In some cases (e.g., copper futures in the single-beta model), the mean adjusted R^2 from the rolling regressions substantially exceeds that from the unconditional regression, indicating significant time-series variation in conditional beta estimates. More typically, the adjusted R^2 is little altered, indicating relatively little time-series variation in the rolling beta estimates.

6. Tests of Futures Market Integration

The hypothesis that futures markets are fully integrated with asset markets is tested against both a general alternative and against a specific alternative provided by Hirshleifer (1988). Since futures positions are settled daily, all hypothesis tests are conducted using daily futures returns.[10]

While tests against a general alternative have some power against all alternatives, power can be low against specific alternatives. More

[10] Though hypothesis tests are conducted using daily returns, the results reported are based on systematic risk estimated using monthly returns. This is necessary in the case of multiple betas on economic aggregates since macroeconomic variables such as inflation and industrial production are reported on a monthly basis. Single-beta tests are also conducted using betas estimated from daily data [with Scholes-Williams (1977) corrections for nonsynchronous trading]; inference is unaffected.

powerful tests of the null against the alternative provided by Hirshleifer (1988) are obtained by including nonsystematic risk and net hedging data.

6.1 Tests against a general alternative

Equation (2) implies a linear futures pricing function that has a zero intercept and slope coefficients on systematic risk equal to slope coefficients in asset markets. The zero intercept implies that futures positions containing no systematic risk earn zero returns on average. Equality of slope coefficients implies a uniform trade-off of systematic risk for expected return across assets and futures.

To estimate risk premiums in asset markets, 20 size-ranked portfolios of NYSE and AMEX equities are formed. The 5 percent of firms with the smallest equity capitalization at the beginning of a year form portfolio 1 for that year; the 5 percent of firms with the largest equity capitalization form portfolio 20, etc. Construction of equity portfolio returns is described in more detail in the Appendix.

To evaluate the uniformity of risk premiums across assets and futures, I adapt the traditional two-stage methodology of Fama and MacBeth (1973). In the first stage, conditional betas are estimated using data for the 60 months before the month containing date t. For the single-beta model, time series of asset and futures returns are regressed on returns to the CRSP value-weighted equity index. For the multiple-beta model, regressors include the six macroeconomic variables described above in addition to CRSP index returns. The second stage involves cross-sectional regressions of the form

$$R_{pt} = \gamma_{0t} + \gamma^*_{0t} d_p + \sum_{i=1}^{n} [\gamma_{it}\hat{\beta}_{ipt} + \gamma^*_{it}\hat{\beta}_{ipt} d_p] + \epsilon_{pt}, \tag{3}$$

where R_{pt} is the period t return on equity portfolio or futures contract p, $\hat{\beta}_{ipt}$ is the estimated conditional beta for equity portfolio or futures contract p with respect to the ith macroeconomic variable, d_p is an indicator variable equal to zero for equity portfolios and equal to unity for futures contracts, an overcaret denotes a parameter estimate, and n equals 1 or 7 for the single-beta and multiple-beta models, respectively.

The proposition that risk premiums are uniform across asset and futures markets is evaluated by testing whether the estimates of γ^*_{it} equal zero. The proposition that the zero-beta rate for futures contracts is zero is evaluated by testing whether estimates of $\gamma^f_{0t} \equiv \gamma_{0t} + \gamma^*_{0t}$ equal zero.

The cross-sectional regression is estimated once for each period. On each date, I use returns for all futures trading on that day. Each

Table 2
Estimates of systematic risk in monthly futures returns (t-statistics in parentheses)

	Single beta		Multiple betas								Macro betas
	VW	Adj. R^2	VW	CEI	UEI	CT3	CTS	CRP	CIP	Adj. R^2	F-stat. (p-value)
Financial futures											
Treasury bills (n = 167)	.008	.028	−.001	−.081	.017	−.328	−.164	.027	.008	.868	177.52
	(2.42)*	[.041]	(−.41)	(−1.67)*	(0.96)	(−17.83)**	(−7.26)**	(0.84)	(0.97)	[.848]	[.000]
Treasury bonds (n = 148)	.276	.094	.006	1.470	.389	−8.454	−8.675	−2.144	.006	.947	392.77
	(4.05)**	[.140]	(0.34)	(2.32)*	(1.61)	(−34.93)**	(−29.00)**	(−.50)	(0.05)	[.954]	[.000]
Treasury notes (n = 91)	.162	.070	−.002	1.270	.488	−6.588	−6.389	−.694	.237	.953	280.38
	(2.80)**	[.045]	(−.15)	(2.17)*	(2.15)*	(−27.96)**	(−23.62)**	(−1.71)*	(2.16)*	[.963]	[.000]
Eurodollars (n = 91)	.006	.026	.003	−.036	.007	−.339	−.194	−.003	.010	.777	51.40
	(1.84)*	[−.004]	(1.27)	(−.47)	(.33)	(−11.16)**	(−5.60)**	(−.06)	(0.77)	[.726]	[.000]
S&P 500 Index (n = 91)	1.005	.968	1.024	1.455	.350	.532	.325	1.26	.350	.971	2.66
	(52.10)**	[.974]	(46.11)**	(1.78)*	(1..41)	(1.62)	(0.86)	(2.22)*	(2.29)*	[.978]	[.021]
Value Line Index (n = 91)	1.096	.914	1.126	1.826	−.502	1.253	.727	−.105	.430	.924	2.85
	(31.16)**	[.924]	(27.97)**	(1.23)	(−1.12)	(2.10)*	(1.06)	(−.10)	(1.56)	[.930]	[.014]
Agricultural futures											
Live cattle (n = 275)	.138	.009	.133	−.347	−1.332	−.044	−.550	−1.761	−.147	−.004	0.42
	(1.80)*	[.005]	(1.55)	(−.10)	(−1.05)	(−.03)	(−.36)	(−.77)	(−.30)	[.012]	[.868]
Soybeans (n = 275)	−.251	.013	−.236	4.398	0.567	−.350	−.193	−4.78	.652	.009	0.84
	(−2.17)*	[.042]	(−1.82)*	(0.84)	(0.30)	(−.17)	(−.08)	(−1.39)	(0.87)	[.046]	[.543]
World sugar (n = 275)	.097	−.003	.138	4.203	−1.760	1.341	1.303	1.278	−1.215	−.020	0.25
	(0.48)	[.001]	(0.62)	(0.47)	(−.53)	(0.38)	(0.32)	(0.22)	(−.94)	[.021]	[.960]
Wheat (n = 275)	.010	−.004	.029	4.932	−.151	.106	−.461	−.606	.001	−.019	0.35
	(.10)	[.019]	(0.27)	(1.16)	(−.10)	(.06)	(−.24)	(−.21)	(.01)	[.024]	[.909]
Cotton (n = 275)	−.052	−.003	.047	9.546	1.035	2.418	2.097	−.616	.054	.020	2.05
	(−.59)	[−.005]	(0.48)	(2.42)*	(0.72)	(1.56)	(1.20)	(−.24)	(.10)	[.007]	[.060]
Corn (n = 275)	−.040	−.006	.013	1.146	−1.823	2.288	2.012	−1.215	4.307	−.003	0.89
	(−.42)	[.024]	(0.12)	(0.27)	(−1.16)	(1.36)	(1.05)	(−.04)	(0.70)	[.090]	[.505]
Metals and oil futures											
Gold (n = 179)	.093	−.002	.124	−2.526	.950	.744	2.600	−.234	.439	−.003	0.98
	(.82)	[.060]	(1.00)	(−.56)	(.56)	(.43)	(1.23)	(−.08)	(.57)	[.042]	[.440]

Table 2
Continued

	Single beta		Multiple betas								Macro betas F-stat.
	VW	Adj. R^2	VW	CEI	UEI	CT3	CTS	CRP	CIP	Adj. R^2	(p-value)
Silver (n = 275)	.284	.012	.445	−5.779	1.942	6.710	6.758	4.485	.942	.034	2.02
	(1.97)*	[.011]	(2.83)**	(−.91)	(0.83)	(2.68)**	(2.38)*	(1.08)	(1.04)	[.026]	[.063]
Platinum (n = 263)	.299	.017	.371	4.494	.889	1.776	2.672	−2.278	.236	.012	0.80
	(2.34)*	[.047]	(2.64)**	(.79)	(0.43)	(.80)	(1.05)	(−.61)	(.29)	[.023]	[.573]
Copper (n = 275)	.221	.013	.185	−5.017	−.949	−1.188	−1.521	−6.564	.219	.015	1.12
	(2.03)*	[.088]	(1.54)	(.351)	(−.53)	(−.61)	(−.70)	(−2.06)*	(.31)	[.069]	[.352]
Crude oil (n = 81)	−.026	−.012	.310	−9.208	−1.981	4.152	8.314	−.540	3.467	.033	1.62
	(−.10)	[−.019]	(1.01)	(−.79)	(−.56)	(.89)	(1.54)	(−.07)	(1.61)	[.040]	[.154]
Foreign currency futures											
Canadian dollar (n = 211)	.058	.042	.041	−.613	−.243	−5.840	−3.942	−6.354	−2.712	.092	2.93
	(3.20)**	[.050]	(2.09)*	(−.84)	(−.90)	(−2.01)*	(−1.15)	(−1.31)	(−2.37)*	[.111]	[.009]
British pound (n = 211)	.011	−.004	−.006	.727	−.904	−.663	−.104	.929	−.253	.017	1.79
	(.23)	[.006]	(−.11)	(.36)	(−1.22)	(−.84)	(−.11)	(.70)	(.81)	[.001]	[.108]
Japanese yen (n = 211)	−.025	−.004	−.070	.601	−1.246	−1.790	−1.295	−1.663	.215	.021	1.86
	(−.50)	[−.003]	(−1.26)	(.29)	(−1.61)	(−2.15)*	(−1.32)	(−1.20)	(.66)	[.021]	[.090]
German mark (n = 211)	.014	−.004	−.010	3.499	−1.861	−1.076	−.576	−.123	−.232	.042	2.71
	(.28)	[.008]	(−.18)	(1.66)*	(−2.39)*	(−1.29)	(−.59)	(−.09)	(−.71)	[.054]	[.015]
Swiss franc (n = 211)	−.010	−.005	−.044	3.547	−2.073	−1.526	−.742	−.294	(−.416)	.067	3.69
	(−.17)	[.005]	(−.73)	(1.54)	(−2.43)*	(−1.67)*	(−.69)	(−.19)	(−1.15)	[.084]	[.002]

Single betas are from simple regressions of monthly futures return on the return to the CRSP value-weighted index. Multiple betas are from multiple regressions of monthly futures returns on (1) CRSP value-weighted index (VW), (2) change in expected inflation (CEI), (3) unexpected inflation (UEI), (4) change in the three-month Treasury-bill yield (CT3), (5) change in the interest rate term structure (T-bond yield − T-bill yield, CTS), (6) change in risk premium (BAA bond yield − T-bond yield, CRP); and (7) unexpected change in industrial production (CIP). n denotes the sample size in months. The F-statistic is for the hypothesis $B_2 = B_3 = B_4 = B_5 = B_6 = B_7 = 0$. Each coefficient estimate is computed using all available data. For comparison, square brackets denote the time-series mean of adjusted R^2 obtained from 60-month rolling regressions.

* Significantly different from zero at $p < .10$.

** Significantly different from zero at $p < .01$.

regression is estimated using weighted least squares (WLS), with weights equal to the inverse of first-stage residual variances. This procedure yields a time series of estimates on each coefficient. The final coefficient estimate is the mean of the daily estimates.[11]

Results of estimating (3) for the single-beta model are reported in the left columns of Table 3. Consistent with theories of asset market equilibrium and prior empirical work, the estimated price of risk in asset markets (γ_1) is significantly positive (9.1 percent per year). The estimated zero-beta rate in asset markets (γ_0) is positive (2.2 percent per year), but not significant.

Equation (2) implies that time-series means of the estimates of γ_{1t}^* and γ_{0t}^f should equal zero. The estimate of the mean intercept in futures markets (γ_0^f) is not significant (1.0 percent per year). The estimate of the mean coefficient for the difference between the beta risk premium in asset markets and the premium in futures markets (γ_1^*) is not statistically significant at conventional levels ($p = .170$), but the point estimate of -10.2 percent per year is substantial.[12] Table 3 also shows results of testing the hypothesis that times-series means of γ_{1t}^* and γ_{0t}^f estimates are jointly zero.[13] The resulting F-statistic is 1.70 ($p = .184$). Thus, the hypothesis that asset and futures markets are integrated is not rejected at conventional confidence levels.[14]

Results of estimating (3) for the more general multiple-beta model are reported in the right columns of Table 3. The integration hypoth-

[11] Shanken (1992) shows that WLS estimation of the daily cross-sectional regressions provides final coefficient estimates that are asymptotically more efficient than those provided by OLS estimation. Shanken also shows that the standard error of the final estimate obtained in the traditional Fama–MacBeth procedure is a downward biased and inconsistent estimate of the true standard error. The source of the bias is the use of estimated rather than true betas. Shanken develops an errors-in-variables correction that yields asymptotically correct standard errors, which is implemented here. I also investigate the importance of the number of futures markets open for trading, and hence the number of returns used in the daily cross-sectional regressions, increasing over time. This may introduce heteroskedasticity into the time series of daily coefficient estimates. To assess sensitivity to heteroskedasticity, I also compute final coefficient estimates by WLS time-series regressions of daily coefficient estimates on a constant, using the number of markets open each day as the weighting variable. As is the case for portfolio means reported in Table 1, t-statistics are little altered and conclusions are unaffected.

[12] Some additional insight is obtained by computing means while excluding estimates obtained during January. The resulting difference between the beta risk premium estimates in assets and futures is reduced to 2.2 percent per year, implying that economically significant differences in risk premiums across assets and futures are confined to the month of January.

[13] The joint hypothesis is tested using Hotelling's T^2 test, with a correction for errors-in-variables developed by Shanken (1992). The T^2-statistic is a weighed sum of squared coefficient estimates, and a transformation of the T^2-statistic is distributed F.

[14] Since (2) predicts that γ_0^f and γ_1^* should always equal zero, estimates of these parameters should be uncorrelated with information known before the estimation date. To test this implication, I regress the time series of parameter estimates obtained from the single-beta model on a variable shown in prior studies [e.g., Fama and French (1988)] to possess forecasting power for security returns: the lagged dividend yield on the CRSP value-weighted equity index. Results indicate that while lagged dividend yields have forecasting power for asset market risk premiums (t-statistic = 3.09), the time series of estimates of the difference between beta price estimates in assets and in futures is not significantly related to prior divided yields (t-statistic = $-.24$).

Table 3
Estimates of systematic risk prices for equities and futures

	Single-beta model		Multiple-beta model	
	Mean	t-statistic	Mean	t-statistic
Equities				
γ_0 [Intercept]	.022	0.38	.109	1.80*
γ_1 [VW]	.091	2.21*	.004	0.06
γ_2 [CEI]			−.121	−0.51
γ_3 [UEI]			−.726	1.40
γ_4 [CT3]			−2.331	−1.79*
γ_5 [CTS]			2.150	1.78*
γ_6 [CRP]			.080	0.20
γ_7 [CIP]			−1.261	−0.83
Futures				
γ_0^f [Intercept]	.010	1.32	.018	1.94*
γ_1^* [VW]	−.102	−1.37	.221	1.33
γ_2^* [CEI]			.131	0.39
γ_3^* [UEI]			1.858	1.55
γ_4^* [CT3]			3.853	1.26
γ_5^* [CTS]			−4.309	−1.30
γ_6^* [CRP]			.076	0.12
γ_7^* [CIP]			3.378	1.38
F-statistic[1] (p-value)	1.70	(.184)	1.49	(.153)

Coefficients are (250 times) time-series means from estimating

$$R_{pt} = \gamma_{0t} + \gamma_{0t}^* d_p + \sum_{i=1}^{n} [\gamma_{it} \hat{\beta}_{ipt} + \gamma_{it}^* \hat{\beta}_{ipt} d_p] + \epsilon_{pt}$$

for each trading day from January 1967 to December 1989. R_{pt} denotes return in futures or equity market p on day t. d_p is an indicator variable equal to one for futures and zero for equities. β_{ipt} is the coefficient on macroeconomic variable i obtained in a regression of return p on macroeconomic variables over the prior 60 months. $\gamma_0^f \equiv \gamma_0 + \gamma_0^*$ is the estimated intercept for futures. The seven macroeconomic variables are (1) return on the CRSP value-weighted index (VW), (2) change in expected inflation (CEI), (3) unexpected inflation (UEI), (4) change in the three-month Treasury-bill yield (CT3), (5) change in the interest rate term structure (T-bond yield − T-bill yield, CTS), (6) change in risk premium (BAA-bond yield − T-bond yield, CRP), and (7) unexpected change in industrial production (CIP).

[1] The F-statistic is for the hypothesis that the futures coefficients are jointly zero, and is computed using Shanken's (1992) generalization of Hotelling's T^2 test.

* Significantly different from zero at $p < .10$.

** Significantly different from zero at $p < .01$.

esis can be tested by assessing whether the futures intercept (γ_0^f) and the seven slope dummy coefficients (γ_1^* to γ_7^*) equal zero. Though the estimated futures intercept is fairly small in economic terms (1.8 percent per year), the hypothesis that the mean intercept is zero can be statistically rejected ($p = .051$). However, the F-statistic for the joint hypothesis that the intercept and all seven slope dummies are zero again indicates that the hypothesis of integration is not rejected ($p = .153$).

In summary, tests of Equation (2) against a general alternative do not provide sufficient evidence to reject the integration hypothesis. As implied by (2), estimates of the zero-beta rate for futures are near zero, and estimated premiums for systematic risk do not differ significantly across equities and futures. However, these tests may not be very powerful. For example, although the estimate of the systematic risk premium for futures obtained in the single-beta model does not differ significantly from the estimate for equities, it also does not differ significantly from zero (and the point estimate is negative). The following section reports results of more powerful tests against a specific alternative.

6.2 Tests against the net hedging pressure alternative

Conditional means reported in Table 1 indicate that returns in some futures markets are correlated with hedgers' positions. However, this finding provides little direct evidence on Hirshleifer's (1988) alternative hypothesis, which implies that expected futures returns deviate from the linear relation (2) as a function of the quantity of nonsystematic risk in futures market. With changes in notation, the equilibrium relation developed by Hirshleifer (his proposition 3) is

$$E_{t-1}(R^f_{jt}) = \beta^f_{jt-1}\gamma_{1t} \pm \sigma_{jt}\gamma_{2t}, \tag{4}$$

where σ_{jt} is the standard deviation of residual risk for futures market j and γ_{2t} is a (possibly market-specific) parameter that depends on risk aversion and on the magnitude of impediments to futures trading. The sign of the deviation from the integrated market pricing function given by (2) is determined by net hedging pressure in the market; in periods when hedgers hold a net short position, implying that speculators are net long, the expected premium for nonsystematic risk is positive, and vice versa.

To test against this alternative, the product of a net hedging indicator variable and a measure of residual risk is included in the cross-sectional regressions. In months where speculators' net position in market p is long at both the beginning and end of the month containing t, the hedging indicator variable, S_{pt}, is set to 1. In months where speculators are short at the beginning and end of the month, the hedging indicator is set to -1. In months when speculators' net position reverses from beginning to end of month and in months when hedging data are unavailable, the hedging indicator is set equal to 0.

Hirshleifer's model incorporates two market imperfections: fixed setup costs to trading in futures and the inability of producers of the underlying good to sell equity claims. Setup costs are likely to vary across markets, and the nonmarketability of equity claims appears to be most directly applicable to futures on agricultural commodities. Therefore, the empirical specification I use to estimate (4) allows the impact of residual risk to vary across the four groups of futures used in this study.

Let $\hat{\sigma}_{pt}$ denote the standard deviation of the residuals from (rolling 60-month) beta estimation regressions, and let d_j denote an indicator variable set equal to 1 for futures in group j and to 0 otherwise. I use the following cross-sectional specification to examine Hirshleifer's alternative:

$$R_{pt} = \gamma^f_{0t} + \sum_{i=1}^{n} [\gamma^f_{it}\hat{\beta}_{ipt}] + \sum_{j=1}^{4} [\gamma^j_{8t}\hat{\sigma}_{pt}S_{pt}d_j] + \epsilon_{pt}. \qquad (5)$$

Equation (5) is estimated on a daily basis using those futures markets open for trading on each day.[15] The hypothesis that residual risk is not priced in futures markets is examined by testing whether estimates of γ^j_8 equal zero.

Results of estimating (5) for the single-beta and multiple-beta models are reported in the first and second columns of Table 4, respectively. Results indicate that, consistent with (4), residual risk conditioned on net hedging has significant explanatory power for foreign currency and agricultural futures returns. Mean coefficient estimates from the single-beta model are 8.2 percent (annualized return premium per unit of monthly standard deviation) in agricultural commodities and 19.2 percent in foreign currencies, with t-statistics of 3.95 and 5.14, respectively. Results from the multiple-beta model indicate premiums on hedging-conditioned residual risk that are smaller (2.2 percent in agriculturals and 2.7 percent in foreign currencies) than those estimated in the single-beta model, but still significant, with t-statistics of 3.53 and 3.22, respectively. These results imply that, consistent with the implications of Hirshleifer's (1988) analysis, mean agricultural and foreign currency futures returns are greater than the level predicted by (2) in months when speculators are net long, and are less than predicted by (2) in months when speculators are net short, by an amount proportional to residual risk.

Results for metal futures are also consistent with the net hedging pressure alternative, but are weaker. The t-statistic for the estimated premium on hedging-conditioned residual risk in metals is 2.02 in the single-beta model and 1.60 in the multiple-beta model. There is little evidence that hedging pressure affects returns in financial futures: t-statistics on the estimated hedging premium for financials are -0.82 and 0.92 in the single-beta and multiple-beta models, respectively.

To provide a control experiment, I estimate a specification similar to (5) that includes residual risk as an explanatory variable, but does

[15] A version of (5) that includes equity returns and allows for a premium on residual risk in equity markets is also estimated. The estimated coefficient on equity residual risk does not differ significantly from zero (t-statistic = 0.92).

Table 4
Tests of residual risk pricing in futures

	Conditional on net hedging				Not conditional on net hedging			
	Single Beta		Multiple Betas		Single Beta		Multiple Betas	
	Mean	t-statistic	Mean	t-statistic	Mean	t-statistic	Mean	t-statistic
Intercept								
γ^f_0	0.90	0.94	−0.61	−0.48	2.89	0.90	4.77	1.26
Macro betas								
γ^f_1 [VW]	−8.19	−0.85	−3.57	−0.16	−8.67	−0.47	37.01	1.24
γ^f_2 [CEI]			−0.17	−0.34			0.89	1.85*
γ^f_3 [UEI]			20.68	0.49			−0.72	−0.52
γ^f_4 [CT3]			−3.70	−0.81			−13.64	−2.06*
γ^f_5 [CTS]			6.75	1.31			10.06	1.76*
γ^f_6 [CRP]			−1.59	−1.55			2.10	1.77*
γ^f_7 [CIP]			2.11	0.50			11.40	2.51*
Residual risk								
γ_8 (financial)	−9.66	−0.82	1.52	0.92	−6.12	−0.33	2.24	0.23
γ_8 (agricultural)	8.21	3.95**	2.18	3.53**	0.63	0.22	0.36	0.64
γ_8 (foreign currency)	19.23	5.14**	2.71	3.22**	0.74	0.20	−1.53	−1.59
γ_8 (metals)	7.25	2.02*	1.34	1.60	2.59	0.56	−0.76	−0.50

Coefficients are (250 times) time-series means from estimating

$$R_{pt} = \gamma^f_{0t} + \sum_{i=1}^{n} [\gamma^f_{it} \hat{\beta}_{ipt}] + \sum_{j=1}^{4} [\gamma^j_{8t} \hat{\sigma}_{pt} S_{pt} d_j] + \epsilon_{pt}$$

for each trading day from January 1967 to December 1989. R_{pt} denotes the percentage change in future price p on day t. $\hat{\beta}_{ipt}$ is the coefficient on macroeconomic variable i obtained in a regression of return p on macroeconomic variables over the prior 60 months. $\hat{\sigma}_{pt}$ is the standard error of residuals from the beta-estimation regression. For the columns conditional on hedging, S_{pt} is equal to 1 (−1) in months, where speculators are net long (short) in market p at beginning and end of month, and 0 otherwise. For the columns not conditional on hedging, S_{pt} always equals 1. d_j is an indicator variable that equals 1 for futures in group j and 0 otherwise. The seven macroeconomic variables are (1) return on the CRSP value-weighted index (VW), (2) change in expected inflation (CEI), (3) unexpected inflation (UEI), (4) change in the three-month Treasury-bill yield (CT3), (5) change in the interest rate term structure (T-bond yield − T-bill yield, CTS), (6) change in risk premium (BAA-bond yield − T-bond yield, CRP), and (7) unexpected change in industrial production (CIP).

* Significantly different from zero at $p < .10$.

** Significantly different from zero at $p < .01$.

not condition on net hedging (the hedging indicator S_{pt} is set equal to 1 for all observations). Results are reported in the third and fourth columns of Table 4. Estimated coefficients on residual risk do not, for any group of futures in either the single-beta or multiple-beta model, differ significantly from zero. This result is important since a finding that residual risk per se has explanatory power for mean futures returns would be inconsistent with existing equilibrium models of futures pricing. In contrast, finding that residual risk conditioned on net hedging has explanatory power in some futures is consistent with Hirshleifer's equilibrium model.

Equilibrium models based on the assumption that all claims are costlessly marketable imply that expected futures returns depend on systematic risk but not hedging pressure. The absence of a significant association between returns and hedging pressure for financial futures is consistent with the imperfections implicit in hedging pressure theories not affecting equilibrium pricing in financial futures. In contrast, documentation of statistically reliable associations between hedging pressure and mean returns in agricultural, foreign currency, and, to a lesser extent, metal futures markets is consistent with imperfections being important for equilibrium in these markets.

The evidence reported here indicates that net hedging pressure has the strongest effect on returns in foreign currency futures. A key imperfection incorporated by Hirshleifer in his models is the inability of producers to market equity claims. Since it is not obvious that this imperfection applies directly to foreign currency markets, this finding suggests an economically important role for hedging pressure in circumstances broader than those Hirshleifer explicitly considers.

There is a growing literature on demand for hedging by corporations. This demand exists even if shareholders can diversify and hedge on personal account. In Smith and Stulz (1985), corporations hedge risks to reduce expected taxes and the likelihood of financial distress. In Bessembinder (1991), corporations hedge to reduce agency costs and improve contracting terms. DeMarzo and Duffie (1991, p. 262) assert that publicly traded firms account for the "bulk of hedging positions in futures and forward markets," and present a model based on information asymmetries that implies increases in firm value due to hedging. Though the implications of corporate hedging for futures market equilibrium have not been formally assessed, it seems reasonable to conjecture that demand for currency hedging by corporations may underlie the empirical findings of this study.

6.3 Covariation between hedging pressure and beta estimates

The conditional beta estimates used in prior sections do not allow for possible comovement between futures' beta coefficients and net hedging. If such an association exists, the resulting errors-in-variables problem could lead to spurious significance in the cross-sectional regressions.

To investigate this possibility, I estimate betas that covary with net hedging. The reported results are based on a simple methodology introduced by Carter, Rausser, and Schmitz (1983), in which the only source of time variation in betas is hedging pressure.[16] A second more general technique based on methods presented in Davidian and Carroll (1987), in which betas vary on the basis of a conditioning information set that includes hedging pressure (along with variables that measure recent movement in and comovement between return and macroeconomic series), is also estimated, with conclusions uniformly unaltered.

Estimated betas are based on the system

$$R_{pt} = \alpha_{0p} + \sum_{i=1}^{n} \beta_{ipt} F_{it} + \epsilon_{pt}, \tag{6}$$

$$\beta_{ipt} = \alpha_{1p} + \alpha_{2ip} S_{pt}, \qquad \text{for } i = 1,\ldots,n, \tag{7}$$

where (6) defines market-model-style regressions of futures returns on macroeconomic factors that accommodate time-varying betas, and (7) states that time variation in betas results from linear relations between conditional betas and the speculator position variable, S_{pt}, defined above. Parameter estimates are obtained by OLS estimation of the equation resulting when (7) is substituted into (6). Conditional beta estimates are fitted values from (7).

Results of estimating (6) and (7) (not reported) provide some evidence of covariation between betas and hedging pressure. For the single-beta model, the hypothesis that α_{2p} equals zero can be rejected for six of the 22 futures markets. Of the six rejections, three indicate positive and three indicate negative comovement between betas and hedging pressure. In contrast, there is little evidence that betas on macroeconomic variables covary with hedging pressure. The hypothesis that coefficients on all six interaction terms (between macroeconomic betas and the hedging pressure variable) are jointly zero is not rejected in any of the 22 futures markets.

In Table 5, I report results of estimating the cross-sectional regression (5) using fitted values from (7) as beta estimates and standard deviations of residuals from (6) as residual risk estimates. The key findings with respect to hedging pressure are robust to this specification. Estimated premiums for residual risk conditional on net hedg-

[16] Carter, Rausser, and Schmitz (1983) do not detect significant interaction between CFTC-reported hedging and estimated commodity futures betas.

Table 5
Tests of residual risk pricing in futures when betas covary with hedging pressure

	Single-beta model		Multiple-beta model	
	Mean	*t*-statistic	Mean	*t*-statistic
Intercept				
γ_0^f	0.93	1.05	−1.77	−1.53
Macro betas				
γ_1^f [VW]	−1.99	−0.24	−11.81	−0.81
γ_2^f [CEI]			−0.39	−1.01
γ_3^f [UEI]			−0.34	−0.30
γ_4^f [CT3]			−3.32	−1.31
γ_5^f [CTS]			4.31	1.40
γ_6^f [CRP]			−1.19	−1.27
γ_7^f [CIP]			1.31	0.48
Residual risk				
γ_8 (financial)	−3.54	−1.40	1.84	0.38
γ_8 (agricultural)	1.51	3.74**	2.14	3.30**
γ_8 (foreign currency)	3.70	4.88**	3.23	3.33**
γ_8 (metals)	1.03	1.49	0.63	0.66

Coefficients are (250 times) time-series means for estimating

$$R_{pt} = \gamma_{0t}^f + \sum_{i=1}^{n} [\gamma_{it}^f \hat{\beta}_{ipt}] + \sum_{j=1}^{4} [\gamma_{8t}^j \hat{\sigma}_{pt} S_{pt} d_j] + \epsilon_{pt}$$

for each trading day from January 1967 to December 1989. R_{pt} denotes the percentage change in future price p on day t. $\hat{\beta}_{ipt}$ is the estimated hedging-conditioned beta on macroeconomic variable i obtained as the fitted value from text equation (7). $\hat{\sigma}_{pt}$ is the standard error of residuals from the beta-estimation regression. S_{pt} is equal to 1 (−1) in months, where speculators are net long (short) in market p at beginning and end of month, and 0 otherwise. d_j is an indicator variable that equals 1 for futures in group j, and 0 otherwise. The seven macroeconomic variables are (1) return on the CRSP value-weighted index (VW), (2) change in expected inflation (CEI), (3) unexpected inflation (UEI), (4) change in the three-month Treasury-bill yield (CT3), (5) change in the interest rate term structure (T-bond yield − T-bill yield, CTS), (6) change in risk premium (BAA-bond yield − T-bond yield, CRP), and (7) unexpected change in industrial production (CIP).

* Significantly different from zero at $p < .10$.

** Significantly different from zero at $p < .01$.

ing remain positive and highly significant in agricultural and foreign currency futures, under both the single-beta and multiple-beta models. Estimated premiums for residual risk conditional on net hedging in financial and metal futures are again insignificant. Thus, while there is some evidence that futures betas covary with hedging pressure, cross-sectional findings with respect to hedging pressure are robust to this source of variation in betas.

7. Conclusions

In this study, I use prices from equity markets and 22 financial, foreign currency, agricultural, and metal futures markets to provide evidence on the roles of systematic risk and hedging pressure in explaining futures risk premiums. With the exception of some financial futures, there is little evidence that unconditional mean futures returns differ from zero. However, there is substantial evidence that mean returns to nonfinancial futures are nonzero when conditioned on net hedging.

I test a simple model of futures premiums. The null hypothesis is that asset and futures markets are integrated such that expected returns on portfolios that include futures positions do not differ from expected returns on asset-only portfolios of equal systematic risk. This hypothesis is tested against a general alternative by examining its implications that the pricing function for futures has a zero intercept and slope coefficients on systematic risk equal to those in asset markets. I also test this hypothesis against a specific alternative presented by Hirshleifer (1988), which implies that the futures pricing function contains a hedging-dependent, and possibly market-specific, premium for residual risk in addition to premiums for systematic risk.

Consistent with integration, estimates of the intercept in the futures pricing function are close to zero, and the hypothesis that premiums for systematic risk in futures are equal to those in asset markets is not rejected. However, these tests may have relatively low power. While the hypothesis that premiums for systematic risk in futures are equal to premiums in assets cannot be rejected, the hypothesis that premiums for systematic risk in futures are equal to zero also cannot be rejected.

Test results indicate that residual risk per se does not have significant explanatory power for futures returns. However, consistent with the equilibrium model presented by Hirshleifer (1988), evidence exists that residual risk conditional on net hedging has explanatory power for foreign currency and agricultural futures returns. The evidence indicates that, conditional on hedgers holding a net short position in a given futures market, mean futures returns in these markets significantly exceed the level predicted on the basis of systematic risk alone, and vice versa.

The estimated effect of hedging-conditioned residual risk on expected returns is strongest for agricultural and foreign currency futures. The existence of significant effects in foreign currency futures is surprising, since Hirshleifer's model, in particular the assumption that claims to hedgers' profits are nonmarketable, does not seem to apply directly. This finding suggests an economically important role for hedging pressure in broader circumstances than those explicitly considered in his model. One possibility is that equilibrium in foreign currency futures markets is affected by a demand for hedging on the part of widely held corporations.

Models based on the assumption that all claims are costlessly marketable imply that futures premiums depend on systematic risk, but not on hedging pressure. The evidence presented here that hedging pressure is a determinant of premiums in some futures markets is consistent with imperfections being important for equilibrium in these markets. Merton (1987, p. 485) states that "financial models based on frictionless markets and complete information are often inadequate to capture the complexity of rationality in action," and he calls for future research that incorporates market imperfections and information costs. The results presented here support Merton's contention that market imperfections are empirically important and that further insights in understanding the behavior of security returns will derive from extending models to incorporate explicitly market imperfections.

Appendix: Data

Futures prices

Daily settlement prices from 22 markets, from January 1967 (or the inception of trading) to December 1989 are obtained from the Columbia University Futures Center and from Data Resources, Inc. The futures markets are associated with the following exchanges:

Chicago Board of Trade
- Treasury bonds
- Treasury notes
- soybeans
- corn
- wheat

Chicago Mercantile Exchange
- Eurodollars
- Treasury bills
- S&P 500 Index
- Japanese yen
- German mark
- Canadian dollar
- Swiss franc
- British pound
- live cattle

Commodity Exchange, Inc.
- gold
- silver
- copper

New York Mercantile Exchange
- crude oil
- platinum

New York Cotton Exchange
cotton #2

Coffee, Sugar, and Cocoa Exchange
sugar #11

Kansas City Board of Trade
Value Line Stock Index

Equity returns
Twenty equity portfolios, which include all firms continuously listed on the CRSP daily tape in a given year, are formed. For each year from 1967 to 1989 all securities are ranked based on their equity market values at the end of the prior year, and sorted into 20 portfolios. The first portfolio includes the 5 percent of firms with the smallest equity capitalization; the 20th portfolio the 5 percent of firms with the largest equity capitalization, etc. To avoid the bias in mean returns described by Roll (1983), portfolio returns are computed based on equal weighing at the beginning of each year and a buy-and-hold strategy within the year. Buy-and-hold portfolio returns for day t, r_{pt}, are obtained from daily asset returns as

$$1 + r_{pt} = \sum_{i=1}^{n} \left[\prod_{k=1}^{t} (1 + r_{ik}) \right] \left(\sum_{i=1}^{n} \left[\prod_{k=1}^{t-1} (1 + r_{ik}) \right] \right)^{-1},$$

where r_{ik} is the return on security i on the kth day of the year, and n is the number of securities in the portfolio. The numerator is (n times) 1 plus the cumulative return on the portfolio to date t, and the denominator is (n times) 1 plus the cumulative return on the portfolio to date $t-1$.

Macroeconomic variables
Monthly data on U.S. industrial production and inflation (based on the CPI) are obtained from the International Financial Statistics data tape. Interest rate data are obtained from Data Resources, Inc. (DRI). These include the yield on three-month Treasury bills, the yield to maturity on 20-year Treasury bonds, and the yield to maturity on long-term corporate bonds rated BAA by Moodys. The term structure variable is (the change in) the Treasury-bond yield less the Treasury-bill yield. The default premium series is (the change in) the BAA-bond yield less the Treasury-bond yield.

The unexpected change in industrial production series is created as the residuals from an ARIMA(3,1,0) model fitted to the raw industrial production data. The inflation series is decomposed into expected and unexpected components using the "naive interest rate model" evaluated by Fama and Gibbons (1984), who find that this simple method provides inflation forecasts as accurate as more complex methods. Expected inflation is determined as the difference between the one-month Treasury-bill yield and the expected real interest rate, while the expected real interest rate is determined as the simple average of ex post realized real Treasury-bill returns over the prior 12 months.

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