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Can Futures Market Data Be Used to Understand the Behavior of Real Interest Rates?

FREDERIC S. MISHKIN*

ABSTRACT

This paper examines whether futures market data can be used to understand the behavior of real interest rates. Several ways of examining the data indicate that futures market data are not particularly informative about real interest rates. Not only does this evidence cast some doubt on results in previous research that make use of futures market data to draw inferences about real interest rates, but it also indicates that future research on real interest rates may need to turn to a different line of attack.

UNDERSTANDING THE BEHAVIOR OF real interest rates is a central issue in financial economics. In previous research on this topic, several approaches have been used to measure real interest rates. In one approach, real interest rates have been calculated by subtracting survey data on inflation expectations, such as the Livingston data, from nominal interest rates.[1] The problem with survey-based measures of real interest rates is that they are only as good as the survey measure of inflation expectations and there may be little incentive for the survey respondents to answer accurately. An even more telling criticism of survey-based measures, often ignored in the literature, is that the behavior of market expectations is driven by economic agents at the margin who are eliminating unexploited profit opportunities. Market expectations are unlikely, therefore, to be well measured by the average expectations of survey respondents.[2]

Because of doubts about survey-based measures of real interest rates, other researchers have used the assumption of rational expectations and ex-post real interest rate data calculated with aggregate price level data to make inferences about real interest rate behavior.[3] There are several difficulties with this approach

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[1] For example, see Gibson (1972), Cargill (1976), Lahiri (1976), Carlson (1977), Levi and Makin (1979), Tanzi (1980) and Wilcox (1983).

[2] See Mishkin (1981a).

[3] For example, see Mishkin (1981b), Fama and Gibbons (1982), Hamilton (1985) and Huizinga and Mishkin (1986).

as well. First is the need for the rational expectations assumption. Although rational expectations is a maintained hypothesis in much current research in monetary and financial economics, there are questions about its validity particularly when there is a shift from one policy regime to another. An additional problem with this approach is that unanticipated inflation appears as a component of the error term in the statistical analysis. As a result, statistical tests may have low power.[4] Another consequence of the presence of unanticipated inflation in the error term is that the variability of real interest rates cannot be examined directly.

Because of the difficulties with the analysis of ex-post real interest rates derived with aggregate price level data, recent research has begun to use futures market data for contracts on individual commodities to examine real interest rate behavior. This paper examines whether futures market data for contracts on individual commodities can be used to understand the behavior of real interest rates. All the evidence consistently casts doubt on the usefulness of futures market data for understanding real interest rate behavior. Not only does this cast some doubt on results in previous research that make use of futures market data to draw inferences about real interest rates,[5] but it also indicates that future research on real interest rates may need to turn to a different line of attack.

The paper begins with a comparison of approaches for measuring real interest rates, Section II examines the usefulness of futures market data in evaluating real interest rate behavior, and Section III concludes.

I. A Comparison of Approaches

The own-commodity real rate for a particular commodity equals the ex-ante return on a one-period bond in terms of that specific commodity. With the existence of a futures market, an investor can lock in this ex-ante real return at time t by selling the commodity j at a price of S_t^j, use the proceeds to purchase a one-period bond with a nominal return (interest rate) of i_t, and then transform the proceeds received at time $t + 1$ back into the commodity at the futures price set at time t, F_t^j. The own-commodity real rate for commodity j is thus defined as

$$rr_t^j = i_t - \ln(F_t^j/S_t^j), \tag{1}$$

where

rr_t^j = the own-commodity real rate at time t for commodity j,
$\ln(F_t^j/S_t^j)$ = the logarithmic basis,
F_t^j = the futures price at time t for delivery of commodity j at time $t + 1$,[6]
S_t^j = the spot price at time t for commodity j.

[4] See Nelson and Schwert (1977) and Mishkin (1981a) for a discussion of this point.

[5] See Cornell and French (1986) and Hamilton (1987), for example.

[6] Strictly speaking, F_t^j in equation (1) should be the price of a forward contract rather than a futures contract. As Black (1976) has pointed out, futures contracts are priced differently from forward contracts. However, this difference should be minor relative to movements of real interest rates. To the extent that prices of forward contracts and futures contracts differ, this is just one more

An attractive feature of the own-commodity real rate rr_t^j defined above is that it is directly observable at time t and does not require any assumption about expectations formation, such as rational expectations, in order to be measured.

To see how own-commodity real rates are linked to the real interest rate, we can note, as in Cornell and French (1986), that if futures markets existed for all commodities in the commodity bundle of the aggregate price index, then the riskless real interest rate for the aggregate economy, rr_t, can be written as a weighted average of all the own-commodity real rates.

$$rr_t = \sum_{j=1}^{n} \omega_j rr_t^j = i_t - \sum_{j=1}^{n} \omega_j \ln(F_t^j/S_t^j), \tag{2}$$

where

rr_t = the (aggregate) real interest rate,
n = the number of commodities in the economy,
ω_j = the expenditure weight for commodity j—i.e., the relative expenditure on commodity j in the futures commodity bundle—where $\sum_{j=1}^{n} \omega_j = 1$.

Using (1) and (2), we can describe the own-commodity real rate for commodity j as follows:

$$\begin{aligned} r_t^j &= rr_t - [\ln(F_t^j/S_t^j) - \sum_{j=1}^{n} \omega_j \ln(F_t^j/S_t^j)] \\ &= rr_t - \varphi_t^j, \end{aligned} \tag{3}$$

where

φ_t^j = the ex-ante relative price movement for commodity j at time t (assuming there is no risk premium embedded in futures prices due to uncertainty about relative price movements).[7]

We can thus think of each own-commodity real rate as a measure of the aggregate real interest rate which is subject to a measurement error of φ_t^j, the ex-ante relative price movement.

The real interest rate usually studied in the literature is the ex-ante real return on a nominally riskless bond, which is defined as

$$rr_t^* = i_t - \pi_t^e, \tag{4}$$

where

rr_t^* = the expected real return at time t of a one-period bond maturing at time $t + 1$,
π_t^e = the expected rate of change (at time t) in the aggregate price level from time t to $t + 1$.

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reason why futures market data may not provide reliable information about the behavior of real interest rates.

[7] Even if a risk premium reflecting uncertainty about relative price movements were present, the basic thrust of the analysis would remain unchanged. However, the φ_t^j term would have a different interpretation because it would include this risk premium.

Note that this real interest rate is not riskless in real terms because there is uncertainty about the inflation rate. It will thus differ from the riskless real interest rate, rr_t, by a risk premium. It is straightforward to calculate estimates of this risk premium using a capital asset pricing (CAPM) framework, and the estimated risk premium is found to be quite small—11 basis points (0.11 percentage points)[8]—while upper and lower bounds for the possible values of this risk premium are ±40 basis points. These CAPM calculations suggest that even if this risk premium is time varying, it and its possible variation is extremely small relative to measured movements in real interest rates which have undergone large swings in recent years (by as much as 1000 basis points from the late 1970s to the early 1980s).[9] Thus it seems plausible that for all practical purposes we can assume that $rr_t = rr_t^*$, and to simplify the discussion, we will treat rr_t and rr_t^*, as identical in the remainder of the paper.

The intuition for why this risk premium may be so small is that the amount of inflation uncertainty is very small relative to uncertainty for the market return. In the 1967–85 sample period examined in this paper, the standard deviation of inflation uncertainty is less than five percent of the standard deviation of the real market return. Thus the beta for a nominally riskless bond must necessarily also be very small, resulting in a small risk premium.

The real interest rate in (4) is not observable because π_t^e is unobservable, but, as discussed in Mishkin (1981a), we can examine the ex-post real rate which is defined as,

$$eprr_t = i_t - \pi_t^e, \tag{5}$$

where

$eprr_t$ = the realized real return on the one-period bond held from time t to $t + 1$,

π_t^e = the change in the aggregate price level from time t to $t + 1$.

The ex-post real rate can be written in terms of the real interest rate as

$$eprr_t = i_t - \pi_t^e - (\pi_t - \pi_t^e) = rr_t - \epsilon_t, \tag{6}$$

where

$$\epsilon_t = \pi_t - \pi_t^e = \text{the forecast error of inflation.}$$

The equation above thus tells us that the ex-post real rate is just equal to the real interest rate plus an error term which is the forecast error of inflation.

Two disadvantages of using ex-post real interest rate data to examine the behavior of real interest rates are readily understandable from the equation above. The presence of the ϵ_t forecast error of inflation term in ex-post real rates means that the researcher cannot directly examine how the variability of the real

[8] This calculation is found in a longer version of this paper, Mishkin (1989). Bodie, Kane and McDonald (1985) reach a very similar conclusion even though they use a model with different assumptions. Their estimate of 15 basis points for this risk premium is very close to my estimate of 11 basis points.

[9] See Huizinga and Mishkin (1986).

interest rate changes over time. Specifically, without knowledge about the variability of inflation forecast errors, information about the variability of the real interest rate cannot be directly extracted from information about the variability of the ex-post real interest rate. The second disadvantage stems from the fact noted in Mishkin (1981a) that the presence of the ϵ_t inflation forecast error term implies that statistical tests using ex-post real rates will have low statistical power. Equation (3) and our discussion of own-commodity real rates indicates, however, that they are also subject to a similar disadvantage. Specifically, for both ex-post real rates and own-commodity real rates there is a signal to noise problem. For own-commodity real rates, the noise is the ex-ante relative price movement, φ_t^j, while for ex-post real rates, the noise is the inflation forecast error, ϵ_t. If the ex-ante relative price movements, φ_t^j, are far greater in magnitude than movements in the real interest rate, then examining own-commodity real rates is unlikely to help us understand the behavior of real interest rates. Furthermore, if the variance of φ_t^j greatly exceeds the variance of ϵ_t, then we are likely to obtain better information about the behavior of real interest rates from using ex-post real rate data than from own-commodity real rate data constructed using futures market data.

Now that we understand the issues relating to the advantages and disadvantages of using futures market data, we can go on to examine what the data on own-commodity actually looks like.

II. An Examination of the Own-Commodity Real Rate Data

The futures market data were obtained from the Center for the Study of Futures Markets, Columbia Business School. The empirical analysis requires a choice of non-financial commodities with equal spacing between contracts that also were traded for a substantial period both before and after October 1979. Precious metals are not included in the study because they are most likely to be subject to cash-and-carry arbitrage which implies that they will contain little information about real interest rates. These criterion lead to the selection of five commodities: live cattle (id #2), live hogs (id #4), soybeans (id #17), frozen orange juice (id #12), and lumber (id #27). The sample period for cattle, hogs and soybeans extends from January or February, 1967 until November or December 1985. Because earlier data was not available, the sample period for orange juice extends from January 1968 until November 1985, while that for lumber is from January 1971 until November 1985.

These commodities have contracts maturing every two months, so that the observation interval and holding period for the own-commodity real rates is two months long. In calculating the logarithmic basis, it is important to make sure that the future and spot prices pertain to the exact same commodity. Thus the futures price for the maturing contract is used as the spot price for the commodity. For example, the logarithmic basis for cattle (at an annual percentage rate) in February, 1967 is calculated as 600 times the log of the April futures price at the closing on January 31, 1967 minus the log of the February futures price at the closing on January 31, 1967. The nominal interest rate is computed as a continuously compounded percent an annual rate from beginning of month prices of U.S. Treasury bills with two months to maturity (obtained from the Center for Research in Security Prices at the University of Chicago). For the February, 1967 own-commodity real rate, the nominal interest rate is 600 times the log of 100 minus the two-month bill price at the closing on January 31, 1967.

Figure 1 contains plots of the own-commodity real rates for the five commodities. The most prominent feature of the own-commodity real rates is their tremendous variability.[10] The own-commodity rates often are outside the range ±50% (annual rate) and the standard deviations for the five commodities are

Cattle	Hogs	Soybeans	Orange Juice	Lumber
25.9%	41.7%	16.1%	28.5%	39.2%.

As a comparison, each plot in Figure 1 also contains a measure of the ex-ante aggregate real rate calculated using the CPI with the procedure found in Huizinga and Mishkin (1986). (The ex-post aggregate real rate is not shown in addition because it would be indistinguishable in the plots from the ex-ante aggregate real rate.) Not only is the variability of the ex-ante aggregate real rate an order of magnitude smaller than those of the own-commodity rates, but this is also the case for the ex-post aggregate real rate: the standard deviation of the ex-ante aggregate real rate is 2.6%, while it is 3.2% for the ex-post aggregate real rate.

As discussed in the previous section, the ex-post aggregate real interest rate differs from the ex-ante aggregate real interest rate by the forecast error of inflation. On the other hand, the own-commodity real rate differs from the ex-ante aggregate real rate by the amount of the ex-ante relative price movements for the specific commodity. Thus which approach gives more accurate inference about the behavior of real interest rates depends on whether the noise represented by the ex-ante relative price movements is greater than the noise attributable to the forecast error of inflation. The standard deviations of the own-commodity real rates and the ex-post aggregate real rate does provide us with information about the relative size of the noises. The upper bound on the standard deviation of the forecast error of inflation is easily shown to equal the standard deviation of the ex-post real interest rate, 3.2%.[11] The lower bounds on the standard deviations of the ex-ante relative price movements can be computed using the fact that the minimum possible variance of the ex-ante relative price movements

[10] Because the high variability of own commodity real rates is so striking, I did check outliers in the data for accuracy by comparing the Center for Futures Markets data with published quotes in the *Wall Street Journal.* In no case did I find a discrepancy between the two data sources. Furthermore, although at first glance it might appear as though the tremendous variability of an own-commodity real rate implies an unexploited profit opportunity, closer inspection of the market suggests that this is unlikely. For example, Bruce Hamilton has pointed out to me that the variability of the own-commodity real rate for hogs (which is the highest of the five commodities) is exactly what might be expected because hogs must be slaughtered at a specific point in their lives in order to get a good price for their meat.

[11] Because the forecast error of inflation is uncorrelated with the ex-ante aggregate real rate under rational expectations, the variance of the ex-post real rate is the sum of the variances of the ex-ante real rate and the forecast error of inflation. Hence the upper bound on the variance of the forecast error of inflation is reached when the variance of inflation is zero: i.e., when the upper bound equals the variance of the ex-post real interest rate.

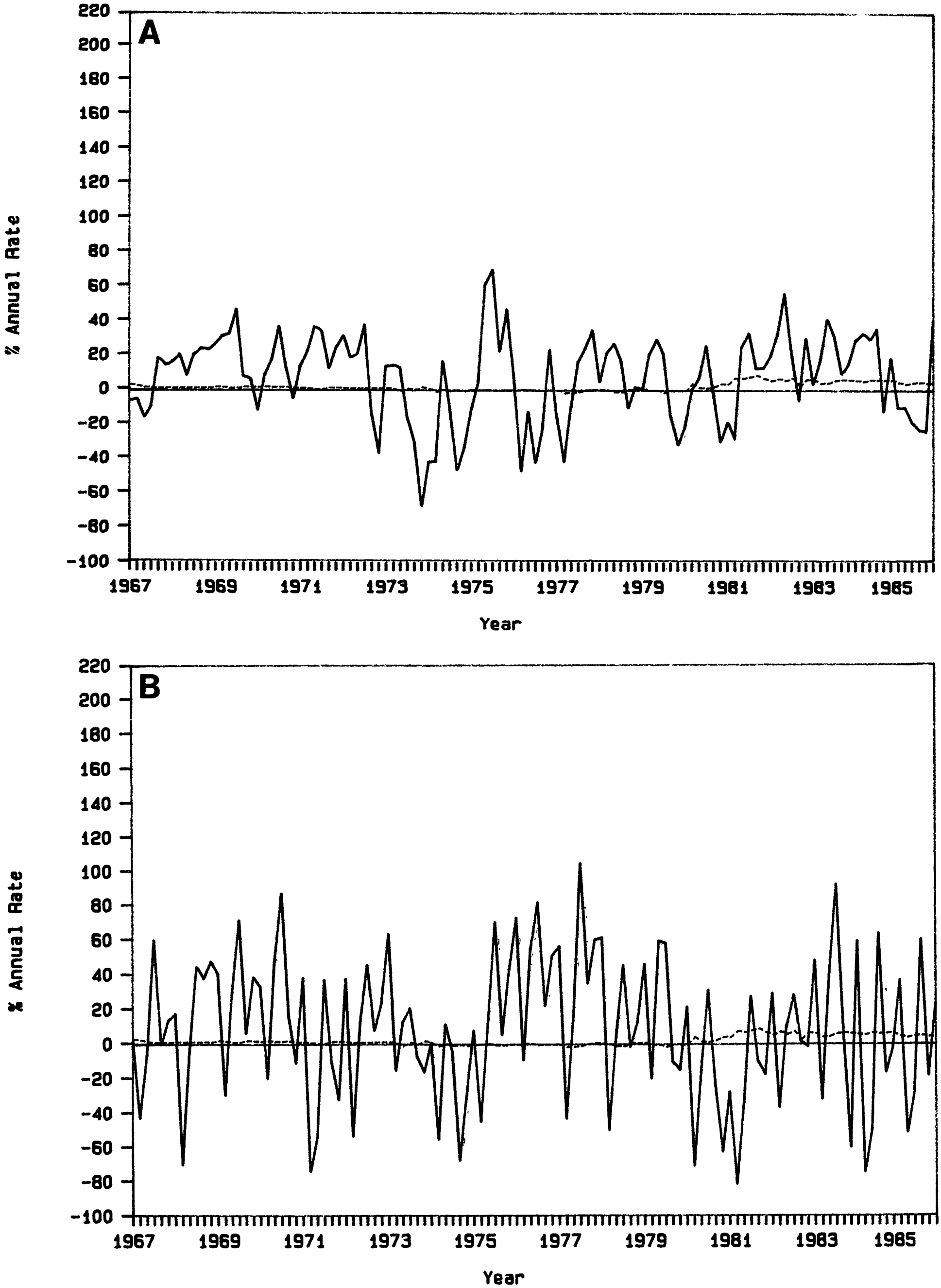
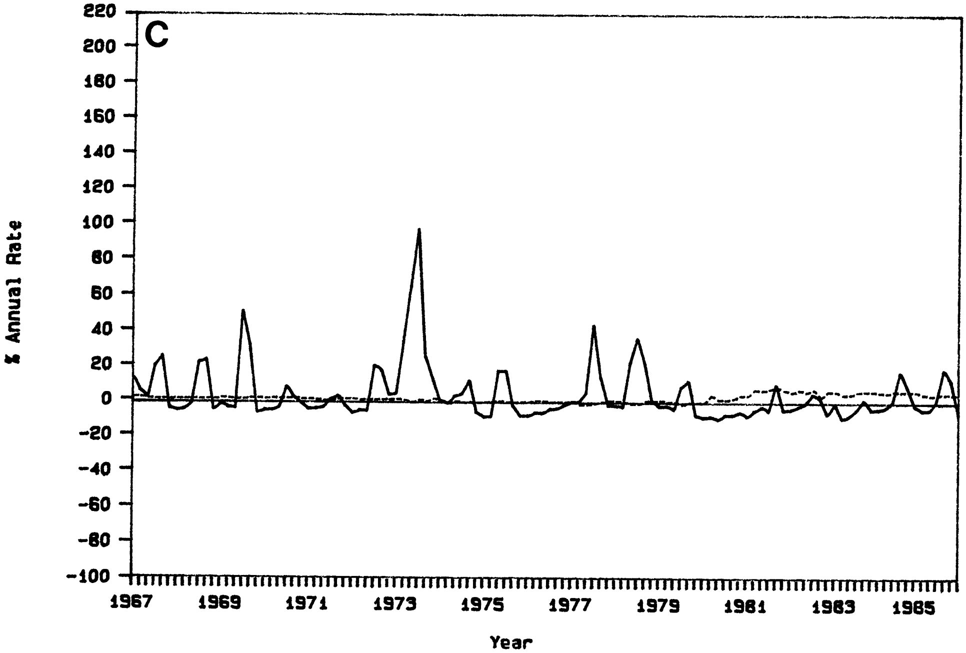
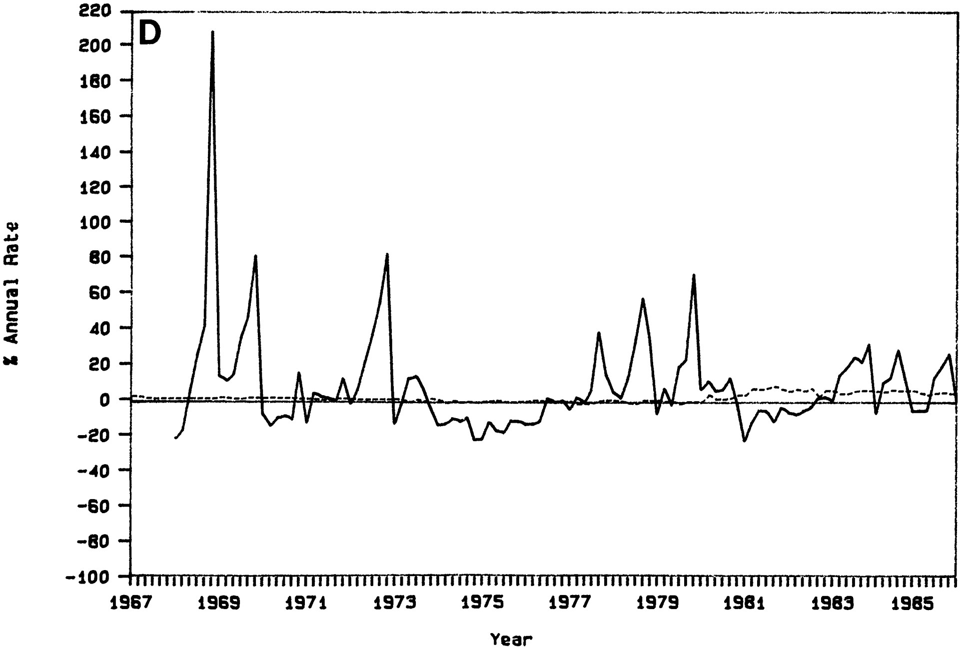


Figure 1. Own-commodity real rates and the aggregate real interest rate. The solid line in each panel is the own-commodity real rate for the commodity named in the panel, while the dashed line is the ex-ante aggregate real interest rate calculated using the CPI with the procedure described in Huizinga and Mishkin (1986). The own-commodity real rate for commodity j is defined as

$$rr_t^j = i_t - \ln(F_t^j/S_t^j),$$

where rr_t^j = the own-commodity real rate at time t for commodity j, $\ln(F_t^j/S_t^j)$ = the logarithmic basis, F_t^j = the futures price at time t for delivery of commodity j at time $t+1$, and S_t^j = the spot price at time t for commodity j. Panel A: Cattle. Panel B: Hogs.





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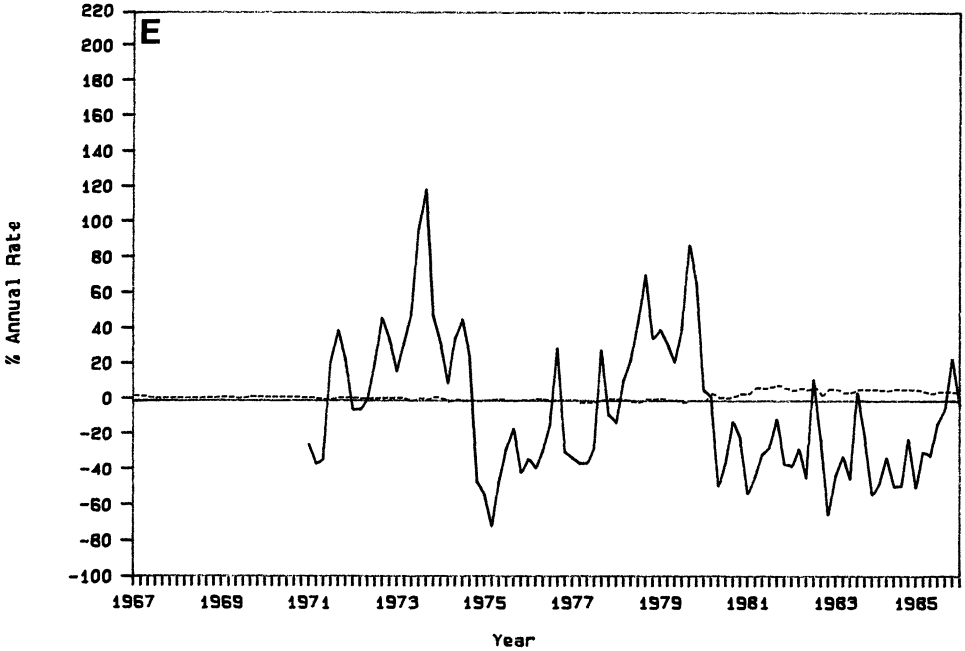

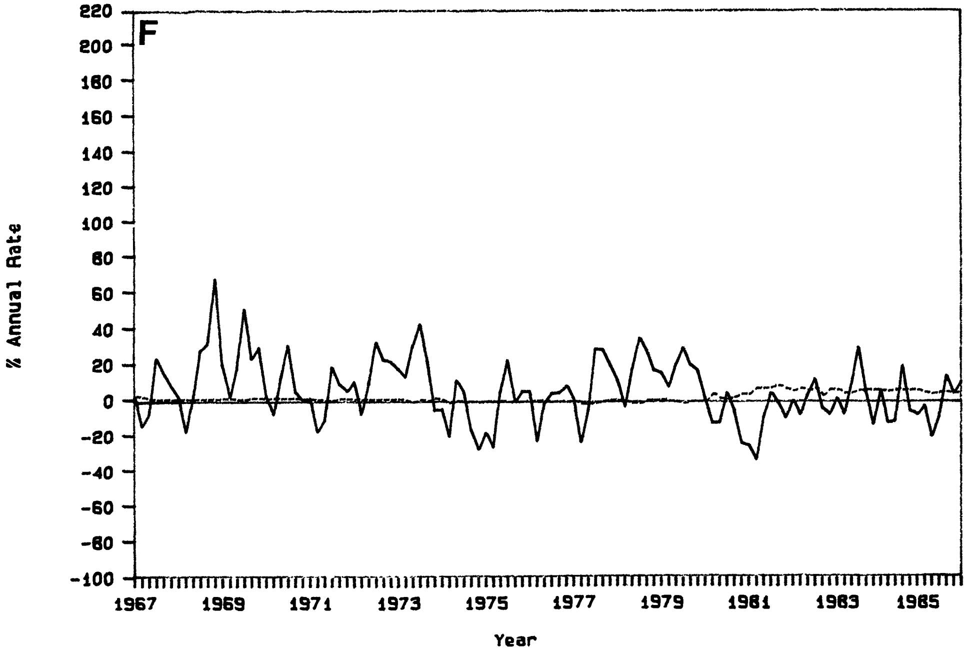

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occurs when the variance of the ex-ante aggregate real rate is at its upper bound and there is perfect negative correlation between the ex-ante relative price movements and the ex-ante aggregate real rate.[12] The lower bounds for the standard deviations of the ex-ante relative price movements for the five commodities are calculated to be the following:

Cattle	Hogs	Soybeans	Orange Juice	Lumber
22.7%	38.6%	13.0%	25.4%	36.0%.

All of these standard deviation are many times greater than the upper bound for the standard deviation of the forecast error of inflation of 3.2%. Hence, the data indicates that the noise in the own-commodity real rates is many times greater than the noise in the ex-post aggregate real rate. Indeed, on average, the variance of the ex-ante relative price movements seems to be on the order of at least 100 times greater than the variance of the forecast errors of inflation. Therefore, futures market data is unlikely to provide as reliable information about the behavior of the ex-ante aggregate real interest rate as ex-post aggregate real interest rate data does.

Since the upper bound for the standard deviation of the ex-ante aggregate real rate equals the standard deviation of the ex-post aggregate real interest rate,[13] 3.2%, the calculations above can be interpreted in a slightly different way. Ex-ante relative price movements (the noise) have a variance that is on the order of over 100 times greater than the variance for the aggregate real interest rate (the signal). Since the signal to noise ratio is apparently less than 1%, own-commodity real rates constructed with futures market data are unlikely to be useful in examining the behavior of aggregate real interest rates.

Another way of examining the value of own-commodity real rate data is to see whether this data confirms two widely accepted facts about the behavior of aggregate real interest rates since October 1979. The first fact is that aggregate real interest rose after October 1979 to levels well above those in the mid to late 1970s. We see evidence of this in Figure 1 where the measure of the ex-ante aggregate real rate averaged −0.8% from the beginning of 1975 until October 1979, while after October 1979 it averaged 4.5%. The conclusion that ex-ante aggregate real interest rates rose sharply after October 1979 is also supported by

[12] The lower bound for the standard deviation of the ex-ante relative price movements equals,

$$\sigma[\varphi^j]_{\min} = \sigma[rr^j] - \sigma[rr]_{\max} = \sigma[rr^j] - 3.2\%,$$

where $\sigma[\varphi^j]_{\min}$ is the lower bound for the standard deviation of the ex-ante relative price movement for commodity j, $\sigma[rr^j]$ is the standard deviation of the own-commodity rate, and $\sigma[rr]_{\max}$ is the upper bound for the standard deviation of the ex-ante aggregate real rate, which, as is shown in the subsequent footnote, equals the standard deviation of the ex-post aggregate real rate, 3.2%. The lower bound, $\sigma[\varphi_j]_{\min}$, is derived using equation (3) to write the variance of the own-commodity rate as

$$\sigma^2[rr^j] = \sigma^2[rr] - 2\rho\sigma[\varphi_j]\sigma[rr] + \sigma^2[\varphi_j],$$

where ρ is the correlation between φ^j and rr. Then recognizing that $\sigma^2[\varphi^j]$ reaches its minimum when $\rho = -1$ and $\sigma[rr]$ is at a maximum, we can solve for $\sigma[\varphi^j]_{\min}$.

[13] Using reasoning similar to that of footnote 11, the upper bound on the variance of the ex-ante real rate is reached when the variance of the forecast error of inflation is assumed to be zero; i.e., when the upper bound equals the variance of the ex-post real rate.

other approaches for measuring the ex-ante aggregate real interest rate, such as those using survey measures of expected inflation. The second fact is the increase in variability of ex-ante aggregate real rates after October 1979: From the beginning of 1975 to October 1979, the standard deviation of the ex-ante real rate measure in Figure 6 is 0.8%, and after October 1979 it is 2.2%. These two facts about the behavior of ex-ante real interest rates seem to be accepted by almost all economists, and yet they are not at all evident in the own-commodity real rate data.

In the first five panels of Figure 1 and the last panel which displays the average of the own-commodity real rates, we see that, rather than rising after October 1979, there was some tendency for the own-commodity real rates to fall: three of the own-commodity real rates (hogs, soybeans and lumber) tend to decline after October 1979, as does the average of all five own-commodity real rates. The average values of the own-commodity real rates for the five commodities in the 1975 to October 1979 period and the post-October 1979 period are shown in Table I. In addition, the variability of the own-commodity real rates usually fall after October 1979 rather than rise, as can also be seen in Table I.[14] Since the behavior of the own-commodity real rates before and after October 1979 seems to be at odds with what most analysis suggests occurred for ex-ante aggregate real rates, we have further evidence that using futures markets to examine real interest rate behavior may not be particularly illuminating.

Table I

Average Levels and Standard Deviations of Own-Commodity Real Rates

The own-commodity real rate for commodity j is thus defined as $rr^j_t = i_t - \ln(F^j_t/S^j_t)$, where rr^j_t = the own-commodity real rate at time t for commodity j, $\ln(F^j_t/S^j_t)$ = the logarithmic basis, F^j_t = the futures price at time t for delivery of commodity j at time $t + 1$, and S^j_t = the spot price at time t for commodity j.

Panel A. Average Levels of Own-Commodity Real Rates

	Cattle	Hogs	Soybeans	Orange Juice	Lumber	Average of Own-Rates
1975 to October 1979	7.7%	27.4%	4.1%	3.8%	−3.4%	7.9%
post-October 1979	9.0%	−4.2%	−2.1%	6.7%	−25.3%	−3.2%

Panel B. Standard Deviations of the Own-Commodity Real Rates

	Cattle	Hogs	Soybeans	Orange Juice	Lumber	Average of Own-Rates
1975 to October 1979	28.8%	39.8%	13.8%	18.9%	40.1%	16.6%
post-October 1979	24.1%	42.7%	7.2%	16.9%	25.8%	13.1%

[14] The finding here is however consistent with the evidence in Parks (1978) who suggests that there is a positive correlation between the level of inflation and relative price variability. Since the inflation level fell after October 1979, relative price variability would be expected to fall and this would be reflected in lower standard deviations of own-commodity real rates.

III. Conclusions

Rarely in monetary/macro economics does empirical work pay sufficient attention to the quality of the data it is analyzing. This paper is an attempt to be an exception since it asks whether a particular data set is appropriate for answering an important set of questions. The analysis here explores several pieces of evidence which provide information on the usefulness of own-commodity real rates constructed from futures market data for understanding the behavior of real interest rates. The evidence can be summarized as follows:

1. The noise in own-commodity real rates (the ex-ante relative price movements) is so large relative to the signal (the aggregate real interest rate) that own-commodity real rates are unlikely to contain much information about aggregate real interest rates.
2. Data on own-commodity real rates is not consistent with several widely accepted facts about the behavior of aggregate real interest rates, indicating that own-commodity real rates can provide misleading information about aggregate real interest rates.[15]

This paper would have a happier ending if it concluded that futures market data is useful for understanding real interest rate behavior. Instead, it indicates that a promising research line exploiting futures market data for analyzing real interest rates is not so promising. In order to learn more about real interest rates, we must pursue a different research route.

[15] In a longer version of this paper, Mishkin (1989), econometric analysis of own-commodity real rate behavior fails to find a shift in the stochastic process which has been documented for aggregate real interest rates. The failure of own-commodity real rate data to reveal a shift in the real rate process provides further evidence that any advantages of own-commodity real rate data for econometric analysis is overcome by the disadvantages stemming from the fact that the noise in the data is orders of magnitude larger than the signal.

REFERENCES

Black, Fischer, 1976, The pricing of commodity contracts, *Journal of Financial Economics* 3, 167–179.

Bodie, Zvi, Alex Kane, and Robert McDonald, 1985, Inflation and the role of bonds in investor portfolios, in Benjamin Friedman, ed.: *Corporate Capital Structures in the United States* (University of Chicago Press, Chicago, IL).

Cargill, Thomas F., 1976, Anticipated price changes and nominal interest rates in the 1950s, *Review of Economics and Statistics* 58, 364–367.

Carlson, John A., 1977, Short-term interest rates as predictors of inflation: Comment, *American Economic Review* 67, 469–475.

Cornell, Bradford and Kenneth French, 1986, Commodity own rates, real interest rates, and money supply announcements, *Journal of Monetary Economics* 18, 3–20.

Fama, Eugene and Michael R. Gibbons, 1982, Inflation, real returns and capital investment, *Journal of Monetary Economics* 14, 297–323.

Gibson, William E., 1972, Interest rates and inflationary expectations: New evidence, *American Economic Review* 62, 854–865.

Hamilton, James, 1985, Uncovering financial markets expectations of inflation, *Journal of Political Economy* 93, 1224–1241.

———, 1987, Monetary factors in the Great Depression, *Journal of Monetary Economics* 19, 145–169.

Huizinga, John and Frederic S. Mishkin, 1986, Monetary policy regime shifts and the unusual behavior of real interest rates, *Carnegie-Rochester Conference Series on Public Policy* 20, 231–274.

Lahiri, Kajal, 1976, Inflationary expectations: Their formation and interest rate effects, *American Economic Review* 66, 124–131.

Levi, Michael D. and John H. Makin, 1979, Fisher, Phillips, Friedman and the measured impact of inflation on interest, *Journal of Finance* 34, 35–52.

Mishkin, Frederic S., 1981a, The real interest rate: An empirical investigation, *Carnegie-Rochester Conference Series on Public Policy* 15, 151–200.

———, 1981b, Are market forecasts rational, *American Economic Review* 71, 295–306.

———, 1989, Can futures market data be used to understand the behavior of real interest rates, First Boston Working Paper Series #FB-87-18R, Graduate School of Business, Columbia University.

Nelson, Charles R. and G. William Schwert, 1977, Short-term interest rates as predictors of inflation: On testing the hypothesis that the real rate is constant, *American Economic Review* 67, 478–486.

Parks, Richard W., 1978, Inflation and relative price variability, *Journal of Political Economy* 86, 79–96.

Tanzi, Vito, 1980, Inflationary expectations, economic activity, taxes and interest rates, *American Economic Review* 70, 12–21.

Wilcox, James A., 1983, Why real interest rates were so low in the 1970s, *American Economic Review* 73, 44–53.