Research Program in Finance
WORKING PAPER SERIES

FINANCE WORKING PAPER NO. 1

OBSERVED PRICE EXPECTATIONS AND INTEREST RATES

By
David H. Pyle

Research Program in Finance Working Papers are preliminary in nature; their purpose is to stimulate discussion and comment. Therefore, they should not be cited or quoted in any publication without the permission of the author. Single copies of a paper may be requested from the Institute of Business and Economic Research.
RESEARCH PROGRAM IN FINANCE AT THE
WALTER A. HAAS SCHOOL OF BUSINESS,
UNIVERSITY OF CALIFORNIA, BERKELEY

The Research Program in Finance in the Walter A. Haas School of Business at the University of California has as its purpose the conduct and encouragement of research in finance, investments, banking, securities markets, and financial institutions. The present reprint and working paper series were established in 1971 in conjunction with a grant from the Dean Witter Foundation.

INSTITUTE OF BUSINESS AND ECONOMIC RESEARCH
Carl Shapiro, Director

The Institute of Business and Economic Research is an organized research unit at the University of California, Berkeley, whose mission is to promote research by faculty and graduate students in the fields of business and economics. The Institute carries out its mission by organizing programs and activities that enrich the research environment, administering extramural research awards, publishing working papers, and making direct grants for research.
OBSERVED PRICE EXPECTATIONS AND INTEREST RATES

David H. Pyle *

This paper is a report on the use of observed price expectations in estimating the effect of expected price changes on nominal interest rates (i.e., the Fisher effect (1896, 1930)). Recent estimates of the Fisher effect, including those by Yohe and Karnosky (1969) and Feldstein and Eckstein (1970), have followed Fisher in using distributed lags on past prices as proxies for price change expectations. Tests of the Fisher effect based on such estimates are joint tests of the distributed lag proxy for price expectations and the effects of price expectations themselves. The use of observed price expectations allows an estimate of the Fisher effect which is not dependent on the assumption that price expectations are based on past prices.

The results reported here show that observed price expectations are at least as powerful as distributed lag proxies in explaining nominal interest rates. Furthermore, there are important differences between the estimates of the Fisher effect obtained by using observed price expectations and the estimates obtained by using distributed lag proxies.

The Data

Data on expected price changes are based on information obtained in a semi-annual survey conducted by Joseph A. Livingston of the Philadelphia Bulletin. Between forty and sixty business economists, including bankers, financiers, and other privately employed economists, government economists, union economists, and academic economists from all over the United States, are asked to provide two predictions (one for six months ahead and one for twelve months ahead) for a number of economic series and indicators. The averages of these forecasts have been published in the Bulletin in January and July of each year since the late 1940’s. The Consumer Price Index is among the series forecast, and it was used by Turnovsky (1970) to generate the six-month and twelve-month rate of price change expectations series which he has kindly provided for use in this study. There is, of course, no guarantee that these observed expectations are suitable measures of the aggregate price change expectations held by individuals who transacted in the bond markets during the period studied. The fact that they are based on the expectations of a cross section of informed individuals in the business world is perhaps reassuring in this regard. Their major virtue for this study is that they constitute a series of expected price changes which is not directly dependent on a distributed lag on past price changes.

Since only six-month and one-year price change forecasts are available using the Livingston data, this study has been restricted to the estimation of Fisher effects for six-month and one-year interest rates. Specifically, yields on six-month finance paper and one-year United States government securities for the period 1954–1969 have been used as the dependent variables in the regressions reported below. The semi-annual yield observations are monthly averages for June and December since these are presumably the periods during which the price expectations reported in early July and early January, respectively, were formed. Some experiments were carried out using average yields for the quarter prior to the date the forecasts were published. Since the results using quarterly averages were substantially the same as those obtained using the June and December data, only the equations using the latter yields are reported. The past price changes are based on the Consumer Price Index.

Received for publication January 10, 1972. Accepted for publication February 27, 1972.

* I wish to thank Stephen Turnovsky, Richard Thunen and member of the Berkeley-Stanford Finance Colloquium for helpful comments. The research in this paper was supported in part by a grant from the Dean Witter Foundation; computer time was provided by the University of California Computer Center.

1 The yields were taken from An Analytical Record of Yields and Yield Spreads, Solomon Brothers, and adjusted to a bond yield basis when necessary. The finance paper series was chosen since (1) it is based on six-month maturity rather than the four- to six-month maturity which is frequently used, and (2) the series is available back to 1954.
The income and monetary data used in estimating price change expectations in an expanded model are discussed under that heading.

**Price Change Expectations Alone**

In his original formulation of the effect of price expectations on interest rates, Fisher expressed the nominal rate of interest \( r_n \) as the sum of a "real" rate of interest \( r_r \) and the expected rate of price change \( \pi_r^e \).\(^2\) For estimating the relationship, he took the "real" rate to be a constant and introduced a distributed lag on past rates of price change \( \rho(r) \) as a proxy for the expected rate of price change or

\[
r_n(t) = a_0 + \sum_{i=0}^{\infty} w(i) \rho(t-i) .
\]

(1)

Recent studies using this approach for estimating the Fisher effect include those by Gibson (1970) and Yohe and Karnosky (1969).

In this study I have followed Yohe and Karnosky (Y-K) in using a polynomial distributed lag\(^3\) on past rates of price change as the proxy expectations variable which is to be compared in this study to the observed expectations variable.\(^4\)

\(^2\) In this study, as in others, including Fisher’s, the product term \( (rr \pi_r^e) \) is neglected. My experiments indicated that this term is, in fact, unimportant.

\(^3\) This is the more general term used by Hall (1967) to describe his equivalent approach using ordinary polynomials rather than the Lagrangian interpolation polynomials used by Almon. The polynomial lag coefficients in this study and their standard errors were obtained using an algorithm developed by Hall and others.

\(^4\) Estimates were also made for the period 1954-1969 in the first difference form used by Gibson with changes in observed price change expectations as the independent variable.

\[
\begin{align*}
\Delta r_n(t) &= .0014 + .21 \Delta \pi_r^e(t) & R^2 &= .13 & S.E.R. &= .0680 & D.W.S. &= 2.36 \\
\Delta r_n(t) &= .0006 + .85 \Delta \pi_r^e(t) & .10 & .0073 & .26 & .0074 & .26 \\
\end{align*}
\]

The coefficient for \( \Delta \pi_r^e \) is significant at the 1 per cent level and the coefficient for \( \Delta \pi_r^e \) at the 5 per cent level. As will be true for all subsequent estimates, the coefficient standard errors are given in parentheses below the coefficient and the following definitions apply:

- \( r_n \) = the six-month finance paper interest rate (annualized)
- \( r_n \) = the one-year United States government securities interest rate
- \( \pi_r^e \) = the observed six-month expected rate of price change (annualized)

Table 1 gives the results for the regression of six-month and one-year interest rates on the observed rate of price change data provided by Turnovsky and on two alternative specifications of the polynomial distributed lag proxy using semi-annual observations of past annual rates of price change. The two polynomial specifications are representative of a number of experiments with different lengths of lag and different degrees of polynomial.\(^5\) Equations 1b and 2b in table 1 are based on a second-degree polynomial with the coefficient for \( \rho(r(t-n-1)) \) constrained to be zero. This results in a monotonically declining lag structure. Equations 1c and 2c in table 1 use the sixth-degree polynomial favored by Y-K and the coefficients here form a serpentine lag structure similar to that found by Y-K.

The estimating equation using the observed rate of price change may be written as

\[
r_n(t) = a_0 + a_1 \pi_r^e(t) .
\]

(2)

This is a modification of the basic Fisher hypothesis since that hypothesis implies that \( a_1 \) should be one (or, in view of the product term, one plus \( a_0 \)). However, as Feldstein and Eckstein (1970) suggest, there are a number of reasons for believing that \( a_1 \) need not be one, so no constraint was placed on the coefficient for the observed expected rate of price change.\(^6\) If we accept this possibility, it is clear that the weights \( w(i) \) in equation (1) are each the product of a coefficient of the model of price change expectations formation (say, \( w(i) \)) and the appropriate value of \( a_1 \). Only if we are willing

\[
\pi_r^e = \text{the observed one-year expected rate of price change}
\]

\[
\rho = \text{the annual rate of price change}
\]

\[
R^2 = \text{the coefficient of determination}
\]

\[
S.E.R. = \text{the standard error of the regression}
\]

\[
D.W.S. = \text{the Durbin-Watson statistic}
\]

\(^6\) In each case in which it is used in this paper, the polynomial lag distribution reported is the one among distributions with similar shapes for which the coefficients of the transformed lag variables in the least squares regression had the largest t-statistic values. This procedure for choosing among polynomial lag distributions would appear to be a better method than choosing on the basis of \( R^2 \) or because a given polynomial lag best approximates an unconstrained lag. I wish to thank Richard Thunen for bringing this to my attention.

\(^5\) In fact, using a model similar to Tobin's (1969), it can be shown that the theoretical size of \( a_1 \) is a function of the relative substitutability between the bond in question and money and real capital (see Pyle, 1971).
OBSERVED PRICE EXPECTATIONS AND INTEREST RATES

Table 1. — Interest Rates and Expected Price Changes, 1954–1969 (semi-annual observations)

<table>
<thead>
<tr>
<th>1. Six-month Finance Paper Rate</th>
<th>( R^2 )</th>
<th>S.E.R.</th>
<th>D.W.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td>(a) Observed Six-month Expected Rate of Price Change ( \tau_n(t) = .029 + .65 \ w_n(t) )</td>
<td>(.71)</td>
<td>(.0084)</td>
<td>(1.40)</td>
</tr>
<tr>
<td>( (.0019) ) ( (.075) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(b) Polynomial Lag Proxy (2nd-degree w/zero restriction) ( \tau_n(t) = .022 + \sum_{l=0}^{5} w(l) \rho(l) )</td>
<td>(.65)</td>
<td>(.0094)</td>
<td>(1.00)</td>
</tr>
<tr>
<td>( (.0035) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( w(0) = .36 ) ( (.085) ) ( w(1) = .25 ) ( (.035) ) ( w(2) = .16 ) ( (.040) ) ( w(3) = .09 ) ( (.055) ) ( w(4) = .042 ) ( (.055) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \frac{\sum_{i=0}^{5} w(i)}{(.105)} = 0.015 )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(c) Polynomial Lag Proxy (5th-degree w/o zero restriction) ( \tau_n(t) = .019 + \sum_{l=0}^{7} w(l) \rho(l) )</td>
<td>(.74)</td>
<td>(.0090)</td>
<td>(1.31)</td>
</tr>
<tr>
<td>( (.0038) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( w(0) = .59 ) ( (.16) ) ( w(1) = .26 ) ( (.175) ) ( w(2) = .27 ) ( (.18) ) ( w(3) = .055 ) ( (.14) ) ( w(4) = .34 ) ( (.14) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \frac{\sum_{i=0}^{7} w(i)}{(.125)} = 1.04 )</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>1. One-year United States Government Securities Rate</th>
<th>( R^2 )</th>
<th>S.E.R.</th>
<th>D.W.S.</th>
</tr>
</thead>
<tbody>
<tr>
<td>(a) Observed One-year Expected Rate of Price Change ( \tau_n(t) = .019 + 1.11 \ \pi w_n(t) )</td>
<td>(.86)</td>
<td>(.0058)</td>
<td>(1.24)</td>
</tr>
<tr>
<td>( (.0017) ) ( (.080) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(b) Polynomial Lag Proxy (2nd-degree w/zero restriction) ( \tau_n(t) = .019 + \sum_{l=0}^{7} w(l) \rho(l) )</td>
<td>(.61)</td>
<td>(.010)</td>
<td>(.93)</td>
</tr>
<tr>
<td>( (.0039) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( w(0) = .33 ) ( (.065) ) ( w(1) = .245 ) ( (.036) ) ( w(2) = .17 ) ( (.03) ) ( w(3) = .11 ) ( (.037) ) ( w(4) = .065 ) ( (.045) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \frac{\sum_{i=0}^{7} w(i)}{(.022)} = .96 )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(c) Polynomial Lag Proxy (5th-degree w/o zero restriction) ( \tau_n(t) = .017 + \sum_{l=0}^{7} w(l) \rho(l) )</td>
<td>(.70)</td>
<td>(.0095)</td>
<td>(1.21)</td>
</tr>
<tr>
<td>( (.0035) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( w(0) = .36 ) ( (.15) ) ( w(1) = .26 ) ( (.18) ) ( w(2) = .31 ) ( (.20) ) ( w(3) = .13 ) ( (.20) ) ( w(4) = .32 ) ( (.18) )</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \frac{\sum_{i=0}^{7} w(i)}{(.18)} = 1.04 )</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

to make some assumptions about the \( w(i) \) are we able to obtain an estimate for \( a_1 \) when a distributed lag proxy is used. One such assumption which seems to be implicit in a number of studies is that the \( w(i) \) sum to one. In the equations reported in table 1, note that the coefficient for the one-year observed expected rate of price change (1.11) is almost twice as large as that obtained for the six-month expected rate of price change (0.65). On the other hand, there is little variation among the sums of lag coefficients for the polynomial dis-
distributed lag proxies. The implied coefficients of price change expectations \( (a_i) \) (on the assumption that the \( u'(i) \) sum to one) range from 0.915 to 1.04 for the six-month data and 0.96 to 1.04 for the one-year data.\(^7\)

On the basis of all the reported statistics, the observed expected rates of price change perform at least as well as the distributed lag proxies in explaining nominal interest rates. The fact that the coefficients of \( v'_i \) and \( v_{12} \) are 8.67 and 13.88 times their respective standard errors is perhaps particularly impressive.

Price Change Expectations in an Expanded Model

As noted earlier, Fisher's estimating equation \( (1) \) assumes that the nominal rate corrected for the effect of the price change expectation is a constant. Sargent (1969), Feldstein and Eckstein (1970), and others have suggested that there are other determinants of the bond market equilibrium which will influence nominal rates. Sargent used a loanable funds model of the bond market equilibrium along with geometrically declining distributed lags on past rates of price change, while Feldstein and Eckstein (F–E) combined a liquidity preference function with a third-degree polynomial distributed lag on past rates of price change.\(^8\)

The results reported here use the F–E liquidity preference approach and compare the effects of the observed expected rates of price change with a second-degree polynomial distributed lag proxy using past rates of price change.\(^8\) Other

\(^7\) By varying the length of lag and the degree of polynomial (and disregarding the resulting effect on the reliability of the coefficient estimates), one can obtain a range of 0.9 to 1.25 for the sums of coefficients of the lag distribution.

\(^8\) Some tests were made using the Sargent loanable funds approach along with the observed rates of price change with results for the coefficient of \( v'_i \) (0.28 (0.095)) and the coefficient of \( v_{12} \) (0.54 (0.17)) which are qualitatively similar to those obtained using the liquidity preference approach. In particular, it is noteworthy that the coefficient of \( v_{12} \) is almost twice that found for \( v'_i \).

\(^9\) There are some problems in comparing the results obtained for the liquidity preference plus price change expectations equations presented here and those obtained by F–E. First, there is the difference in the number of observations since the F–E study was based on quarterly data. This does not appear to be a severe problem since a regression of \( n_{12} \) on quarterly data gave substantially the same fit as that reported in equation 2a of table 2. The second problem is the use of short-term interest rates rather than long-term corporate bond rates, which F–E suggest as being

lag specifications were tested with minor effects on the results (except, of course, for the shape of the lag distributions). Furthermore, the lags in the reported equations are monotonically (almost linearly) declining as was the case for the polynomial lag chosen by F–E in their study.

The model formulated by F–E may be written as

\[
rm(t) = b_0 + b_1LMB(t) + b_2LGNP(t)
+ b_3u'(t)
\]

where \( LMB \) is the log of real per-capita monetary base (see Federal Reserve Bank of St. Louis (1968)) and \( LGNP \) is the log of real per-capita private gross national product.\(^{10}\) The monetary base is taken as an exogenous variable and, in this study, I have followed F–E in disregarding the simultaneity between interest rates and aggregate demand.

The results are given in table 2. As was the case when price expectations were used alone, the coefficient for \( v_{12} \) (91) is considerably larger than that obtained for \( v'_i \) (32) while the sums of polynomial lag coefficients in the six-month rate equation and the one-year rate equation are identical and imply, for the sum of \( u'(i) \) equal to one, a coefficient of price change expectations equal to .43.\(^{11}\)

Both of the observed expected rate of price change variables perform well in their respective equations. An analysis of variance test of

most appropriate to the liquidity preference approach. The absence of observed expectations of long-term rates of change of prices precluded a direct test of the F–E equations. The substitution of short-term interest rates in the liquidity preference model may account for the generally poorer fits and lower level of significance for the coefficients of monetary and aggregate demand variables in this study as compared to the F–E study. However, all of the coefficients in this study do have the correct signs.

\(^{10}\) The F–E model can be derived in the following manner. The basic hypothesis is

\[
1 + rm(t) = (1 + rm(t))(1 + \epsilon(t))\]

where \( rm(t) \) is the interest rate which equilibrates the bond market if the expected rate of price change is zero. The exponent \( b_3 \) allows the price change expectations to be less than fully accommodated for in the nominal rate (as the coefficient \( a_3 \), did earlier). The liquidity preference function determining \( rm(t) \) may be written as

\[
1 + rm(t) = e^{\theta(MB(t) + GNPNP(t))}
\]

Substituting this into equation (3a) and taking natural logs of both sides gives us equation (3).

\(^{11}\) A range of 0.39–0.55 in the sum of lag coefficients for the two equations was obtained by varying the degree of the polynomial and the length of the lag.
### Table 2. Interest Rates, Liquidity Preference, and Expected Price Changes, 1954–1969

<table>
<thead>
<tr>
<th>(semi-annual observations)</th>
<th>1. Six-month Finance Paper Rate</th>
<th>2. One-year United States Government Securities Rate</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>(a) Liquidity Preference Variables Alone</strong></td>
<td>$r_{m1}(t) = -0.50 - 0.07 LMB(t) + 0.11 LGNP(t)$</td>
<td>$r_{m2}(t) = -0.30 - 0.03 LMB(t) + 0.11 LGNP(t)$</td>
</tr>
<tr>
<td></td>
<td>$R^2$ = 0.76, S.E.R. = 0.0079, D.W.S. = 1.03</td>
<td>$R^2$ = 0.77, S.E.R. = 0.0078, D.W.S. = 0.94</td>
</tr>
<tr>
<td><strong>(b) Liquidity Preference Variables Plus Observed Six-month Expected Rate of Price Change</strong></td>
<td>$r_{m1}(t) = -0.40 - 0.068 LMB(t) + 0.068 LGNP(t) + 0.32 \pi_t(t)$</td>
<td>$r_{m2}(t) = -0.36 - 0.03 LMB(t) + 0.021 LGNP(t) + 0.01 \pi_t(t)$</td>
</tr>
<tr>
<td></td>
<td>$R^2$ = 0.81, S.E.R. = 0.0070, D.W.S. = 1.40</td>
<td>$R^2$ = 0.88, S.E.R. = 0.0057, D.W.S. = 1.34</td>
</tr>
<tr>
<td><strong>(c) Liquidity Preference Variables Plus Polynomial Lag Proxy (2nd-degree w/zero restriction)</strong></td>
<td>$r_{m1}(t) = -0.31 - 0.011 LMB(t) + 0.075 LGNP(t) + \sum_{i=0}^{5} w(i) \pi'(t-i)$</td>
<td>$r_{m2}(t) = -0.16 - 0.036 LMB(t) + 0.080 LGNP(t) + \sum_{i=0}^{5} w(i) \pi'(t-i)$</td>
</tr>
<tr>
<td></td>
<td>$w(0) = 0.18, w(1) = 0.12, w(2) = 0.075, w(3) = 0.016, w(4) = 0.012$</td>
<td>$w(0) = 0.14, w(1) = 0.10, w(2) = 0.075, w(3) = 0.05, w(4) = 0.032$</td>
</tr>
<tr>
<td></td>
<td>$R^2$ = 0.83, S.E.R. = 0.0067, D.W.S. = 1.59</td>
<td>$R^2$ = 0.83, S.E.R. = 0.0069, D.W.S. = 1.41</td>
</tr>
</tbody>
</table>

The contribution of the alternative price expectations variables shows both $\pi_t^a$ ($F = 8.78$ with 1 and 28 d.f.) and the polynomial lag proxy for the six-month rate ($F = 6.44$ with 2 and 27 d.f.) to be significant contributors to the explanation of $r_{m1}$ at a level less than 1 per cent, while for the one-year rate, $\pi_t^a$ ($F = 25.1$ with 1 and 28 d.f.) is significant at less than the 1 per cent level and the polynomial lag proxy for the one-year rate ($F = 4.56$ with 2 and 27 d.f.) is significant at a level between 5 per cent and 1 per cent. It is particularly interesting to note that the introduction of the observed expected rate of price change has a larger effect on the aggregate demand variable of the liquidity preference function than does the introduction of the distributed lag proxy. This suggests that some measure of aggregate demand may also have an influence on price forecasts.12

**Summary and Conclusions**

The chief result of this study is the discovery that the observed price expectations contribute significantly to the explanation of nominal interest rates. This provides a confirmation of the Fisher effect which is not dependent on the use of a distributed lag proxy.

12Turnovsky (1970) recognized this possibility and attempted to introduce a measure of the level of economic activity as an explanatory variable. The indicator he used with some measure of success was the unemployment rate. However, he did not find the results sufficiently convincing to accept this modification of the basic extrapolative expectations hypothesis.
Secondly, the results obtained here suggest that the use of distributed lag proxies for expectational variables, while econometrically convenient, may give misleading results with respect to the importance of other independent variables and with respect to the impact of the expectations variables themselves.\(^3\)

Finally, the fact that the coefficient of the six-month observed price change expectations was consistently smaller than that obtained for the one-year observed price expectation suggests that changes in maturity of as little as six months may have an effect on the extent to which price expectations are accommodated in nominal interest rates.

REFERENCES


Feldstein (1970) has recently discussed the problem of specification bias when nominal rates are used in equations calling for "real" rates. Clearly, the corrections for this bias implied by the polynomial lag proxies estimated in this study are quite different from those implied by the coefficients of the observed price change expectations.


