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Stock prices–inflation puzzle and the predictability of stock market returns

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Abstract

This paper considers a new perspective on the relationship between stock prices and inflation, by estimating the common long-term trend in the earning–price ratio and inflation. We find that the transitory deviations from this common trend exhibit substantial out-of-sample forecasting abilities for excess returns at short and intermediate horizons.

Keywords: Time-varying expected returns; Stock return predictability; Stock prices–inflation puzzle

JEL classification: G12; G14; E44; C53

1. Introduction

For more than two decades, empirical studies have documented the negative relationship between real stock prices, as reflected in price–dividend ratio or price–earning ratio, and both expected and realized inflation during the post-World War II period (e.g. Sharpe, 2002; Campbell and Vuolteenaho, 2004, and references herein). However, there is less consensus on what drives it. This negative relationship could reflect: (i) a correlation between inflation and expected real economic growth; (ii) the use of nominal interest rates to discount real cash flows by irrational investors (Modigliani and Cohn, 1979); or (iii) a

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subjective inflation risk premium. The difficulty with the first explanation is that if such a relation exists, then it will concern the expected growth over business cycle horizons (i.e. ranging from one quarter to few years) instead of long-term real cash-flow growth.[1] Moreover, a large literature has documented the poor predictability of real dividend growth and real output growth by the equity ratios (e.g. Campbell, 2003). Thus, the two behavioral hypotheses offer more convincing explanations.

In this article, we focus on the subjective inflation risk premium explanation. We consider a present value model with a conditional time-varying risk premium and estimate the common long-term trend in the earning–price ratio and actual inflation. We investigate the role of the transitory deviations from this common trend for forecasting stock returns (S&P 500). We find that these deviations exhibit substantial out-of-sample forecasting abilities for excess stock returns at short and intermediate horizons.

The paper proceeds as follows: Section 2 presents results of estimating the trend relationship among the earning–price ratio and inflation. Section 3 discusses data used in our forecasting regressions. Section 4 reports out-of-sample predictability test results. Section 5 concludes.

2. Estimating the long-term relationship between stock prices and inflation

In our empirical implementation, we use a modified log linear version of the present value model proposed by Campbell and Shiller (1988). Following Sharpe (2002), we decompose the log dividends per share into the sum of the log earnings per share and the payout ratio.[2] Then, the Campbell–Shiller formula can be rewritten as:

$$e_t – p_t = -\frac{\kappa}{1-\rho} + E_t\left[\sum_{j=0}^{\infty}\rho^j r_{t+j} - \sum_{j=0}^{\infty}\rho^j \Delta e_{t+j} - (1-\rho)\sum_{j=0}^{\infty}\rho^j\left(d_{t+j} - e_{t+j}\right)\right], \quad (1)$$

where E_t denotes investors expectations taken at time t, $e_t – p_t$ denotes the log earning–price ratio, r_{t+j} denotes log stock return during period $t+j$, Δe_{t+j} denotes real earning growth in $t+j$, and $d_{t+j} – e_{t+j}$ denotes the log of the payout ratio (dividends/earnings) in $t+j$. The expected return equals the real risk-free interest rate plus a risk premium. ρ and κ are parameters of linearization.

We assume a time-varying risk premium which can be expressed as a linear function of inflation, π_t. This model can be seen as a conditional factor model in the tradition of Cochrane (1996) in which factors at time $t+1$ are scaled by information variables at time t.

Both $e_t – p_t$ and π_t are very persistent series for which we cannot reject a unit root.[3] Consequently, we investigate the co-integration relationship between them. Then, this presumed co-integrating relationship implies that a deviation from the long-run equilibrium impacts positively or negatively the (log) earning–price ratio such that the equilibrium is restored.

Therefore, we test for co-integration using the multivariate trace statistic developed by Phillips and Ouliaris (1990) and the Johansen and Juselius (1992) approach (Trace Test). Table 1 displays test

[1] Empirical studies generally failed to establish any meaningful relationship between inflation and long-term economic growth (see e.g., Bruno and Easterly, 1998).

[2] This reformulation enables us to focus on earnings which are more closely related to economic fundamentals than dividends since they can be affected by shifts in corporate financial policy (see Campbell, 2000).

[3] Based on ADF and KPSS tests. Test results are available upon request.

Table 1
Phillips–Ouliaris and Johansen co-integration tests

Z_t	P–O trace test			Johansen trace test						
	Test stat.	90% CV	95% CV	# of co-int. relation	Test stat.	90% CV	95% CV	Lags	LM(1)	LM(4)
$[ep_t, \pi_t]$	**60.05**	47.59	55.22	0	**23.29**	17.79	19.99	6	5.19	7.78
				1	3.98	7.50	9.13		$(p=0.27)$	$(p=0.10)$

The table reports tests of the null hypothesis of no co-integrating relationships against the alternative of one or more co-integrating vectors. The only deterministic components in the models are the intercept in the co-integration space. "Lags" gives the number of lags in the estimated VAR model. The appropriate lag-length is selected in order to accept the assumption that residuals are white noise based on LM(1) and LM(4) criteria. A test statistic greater than the specified critical value suggests rejection of the null of no co-integration. Significant coefficients at the 5% level are highlighted in bold face.

results.[4] There is sufficient evidence for one non-zero co-integrating vector between e_t–p_t and π_t. The next step in our analysis is to investigate the role of these transitory movements in the earning–price ratio in forecasting stock returns. Before that, it is necessary to obtain consistent estimates of the parameters of the shared trend in log earning–price ratio and inflation. We use the dynamic ordinary least squares (DOLS) developed by Stock and Watson (1993) to estimate the co-integration parameters to eliminate the effects of regressor endogeneity on the distribution of the least squares estimator. Eq. (2) reports the DOLS estimates (ignoring coefficient estimates on the first differences) for the parameters of the shared trend among earning–price ratio and inflation from the fourth quarter of 1951 to the second quarter of 2003:

$$e_t - p_t = \underset{(-35.67)}{-3.11} + \underset{(6.59)}{10.00}\,\pi_t, \tag{2}$$

where the corrected t-statistics appear in parentheses below the coefficient estimates. The estimated co-integrating coefficients suggest that a one percentage point decrease in actual inflation is associated with a 10% decline in the earning–price ratio and thus in real stock prices.[5] We denote, $e\hat{p}i_t$, the deviation of (log) earning–price ratio from its predicted value based on the co-integrating regression (2).

3. Asset returns data

The data set consists of quarterly observations from 1951:Q4 to 2003:Q2. Stock prices, dividends per share, and earnings per share all correspond to the Standard and Poor's Composite Index.[6] Real data are deflated by the Consumer Price Index (All Urban Consumers) published by the BLS. Let r_t denote the real return on the S&P index. The 3-month T-bill rate[7] is used to construct the real return on the risk-free rate, $r_{f,t}$, and the log excess return $(r_t - r_{f,t})$.

[4] Data are described in the next section.

[5] We estimated a vector error correction model for real stock prices, real earnings and inflation. Results suggest that deviations from the shared trend in log earning–price ratio and inflation are better described as transitory movements in real stock prices than as transitory movements in real earnings or inflation.

[6] The S&P 500 data are available from Robert Shiller's home page at http://www.econ.yale.edu/~shiller. The complete documentation for the data sources is also provided here. Data are updated from the Standard and Poor's web site (S&P 500 Earnings and Estimate Report).

[7] Interest rate data come from the FRED II database.

Log price, p_t, is the natural logarithm of the real S&P price level in quarter t. Log dividends, d_t, are the natural logarithm of real dividends per share in quarter t. Log earnings, e_t, are the natural logarithm of real earnings per share in quarter t. Following Lamont (1998), the log dividend payout ratio is d_t–e_t. The stochastically detrended risk-free rate, $rrel_t$, is the T-bill rate minus its last four-quarter average. This relative bill rate is used by Campbell (1991) to forecast stock returns. Following Lettau and Ludvigson (2005), we use the measure of short-term deviations from the long-run co-integration relationship among the natural logarithm of consumption (c), labor income (y) and aggregate wealth (a), henceforth $c\hat{a}y_t$.

4. Out-of-sample predictability test results

This section examines the out-of-sample predictability of excess returns. Some recent studies (e.g. Goyal and Welch, 2003) expressed concern about the apparent predictability of stock returns because while a number of financial variables display significant in-sample predictive ability, they have negligible out-of-sample predictive power. Also, our in-sample forecasting results could suffer from a "look-ahead" bias that arises from the fact that the coefficients used to generate $e\hat{p}i_t$ are estimated using the full sample.

Two cases are considered. First, agents are assumed to know the co-integration parameters of $e\hat{p}i_t$ which are estimated using the full sample. Second, the co-integration parameters are estimated recursively using only information available at the time of forecast. Moreover, we present out-of-sample predictability results using the two-period lagged value of $e\hat{p}i_t$ because this variable is available with a 1-month delay relative to financial indicators.

We use four statistics to compare the out-of-sample performance of our forecasting models: the mean-squared forecasting error (MSE) ratio, the Clark and McCracken's (2001) encompassing test (ENC-

Table 2

One-quarter-ahead forecasts of excess returns: nested comparisons

Row	Comparison unrestricted vs. restricted	MSE_u/MSE_r	ENC-NEW		MSE-F	
			Statistic	99% CV	Statistic	99% CV
Panel A: co-integrating vector reestimated						
1	$e\hat{p}i_t$ vs. AR	0.9392	9.066**	4.251	9.339**	3.970
2	$e\hat{p}i_{t-1}$ vs. AR	0.9472	7.998**	4.251	8.068**	3.970
3	$e\hat{p}i_t$ vs. const	0.9264	11.621**	4.251	11.129**	3.970
4	$e\hat{p}i_{t-1}$ vs. const	0.9344	10.562**	4.251	9.793**	3.970
Panel B: fixed co-integrating vector						
5	$e\hat{p}i_t$ vs. AR	0.9328	12.889**	4.251	10.227**	3.970
6	$e\hat{p}i_{t-1}$ vs. AR	0.9472	10.020**	4.251	7.908**	3.970
7	$e\hat{p}i_t$ vs. const	0.9216	16.266**	4.251	12.004**	3.970
8	$e\hat{p}i_{t-1}$ vs. const	0.9376	13.213**	4.251	9.418**	3.970

The MSE-F statistic is used to test the null hypothesis that the MSE for the restricted model forecasts is less than or equal to the MSE for the unrestricted model forecasts. The ENC-NEW statistic is used to test the null hypothesis that restricted model forecasts encompass the unrestricted model forecasts. We estimate the co-integration parameters recursively in panel A and use the full sample in panel B. A * (**) denotes significance at the 5% (1%) level.

NEW), the McCracken's (2004) equal forecast accuracy test (MSE-F) and the modified Diebold–Mariano (MDM) encompassing test proposed by Harvey et al. (1998). We apply the ENC-NEW and MSE-F tests for nested comparisons and the MDM test for non-nested comparisons. We report the MSE ratio in both nested and non-nested comparisons. The initial estimation period begins with the fourth quarter of 1951 and ends with the first quarter of 1968. The model is recursively reestimated until the end of the sample.

We report results of the out-of-sample one-quarter-ahead nested forecast comparisons of excess returns in Table 2. We consider two restricted (benchmark) models: a model that includes only a constant as a predictor and a model that includes both a constant and the lagged dependent variable as predictive variables. We find that the unrestricted model (which includes $\hat{epi}_t$) has smaller MSE than the constant restricted model or the autoregressive restricted model. Table 2 shows that regardless of whether the co-integrating parameters are reestimated, or whether the one- or two-period lagged value of $\hat{epi}_t$ is used as a predictive variable, both ENC-NEW and MSE-F tests reject the null hypothesis that $\hat{epi}_t$ provides no information about future excess returns at the 1% significance level.

Results of the out-of-sample one-quarter-ahead non-nested forecast comparisons of excess returns are shown in Table 3. We compare alternatively the model 1 in which the lagged value of $\hat{epi}_t$ is the sole predictive variable with "competitor models" in which either the lagged dependent variable, lagged dividend–price ratio, lagged earning–price ratio, lagged dividend payout ratio, lagged detrended bill rate, lagged value of $\hat{cay}_t$ is the sole predictive variable. A constant is included in each of the forecasting equations.

Table 3
One-quarter-ahead forecasts of excess returns: non-nested comparisons

Row	Model 1 vs. model 2	MSE_1/MSE_2	MDM test	
			Test statistic	p-value
Panel A: co-integrating vector reestimated				
1	$\hat{epi}_t$ vs. $r_t - r_{f,t}$	0.963	3.099**	0.002
2	$\hat{epi}_t$ vs. $d_t–p_t^{\blacklozenge}$	0.976	1.972	0.051
3	$\hat{epi}_t$ vs. $e_t–p_t$	0.970	2.740**	0.007
4	$\hat{epi}_t$ vs. $d_t–e_t$	0.948	2.405*	0.018
5	$\hat{epi}_t$ vs. $RREL_t$	0.963	3.003**	0.003
6	$\hat{epi}_t$ vs. $\hat{cay}_t^{\blacklozenge\blacklozenge}$	0.991	1.561	0.121
Panel B: fixed co-integrating vector				
7	$\hat{epi}_t$ vs. $r_t - r_{f,t}$	0.960	3.435**	0.000
8	$\hat{epi}_t$ vs. $d_t–p_t$	0.973	2.660**	0.009
9	$\hat{epi}_t$ vs. $e_t–p_t$	0.967	2.374*	0.019
10	$\hat{epi}_t$ vs. $d_t–e_t$	0.946	2.969**	0.003
11	$\hat{epi}_t$ vs. $RREL_t$	0.960	3.513**	0.000
12	$\hat{epi}_t$ vs. $\hat{cay}_t$	0.996	2.982**	0.003

The MDM test, is a modified Diebold and Mariano (1995) test statistic to test for forecast encompassing between two non-nested models and to account for finite-sample biases. The null hypothesis is that the model 2 encompasses model 1. We estimate the co-integration parameters recursively in panel A and using the full sample in panel B. A * (**) denotes significance at the 5% (1%) level. ◆The inverse encompassing test that under the null hypothesis model 1 encompasses model 2 is not rejected (p-value = 0.675). ◆◆The inverse encompassing test that under the null hypothesis model 1 encompasses model 2 is not rejected (p-value = 0.353).

Table 4
Long-horizon forecasts of excess returns: nested models

k	1	2	3	4	8	12	16	20	24
Panel A: reestimated $e\hat{p}i_t$ vs. C									
MSE_u/MSE_r	0.941	0.898	0.859	0.842	0.850	0.783	0.764	0.818	0.846
ENC-NEW	**8.992**	**14.497**	**17.997**	**20.160**	**16.080**	**24.555**	**31.709**	27.279	25.408
(p-value)	(0.000)	(0.004)	(0.004)	(0.008)	(0.048)	(0.040)	(0.036)	(0.053)	(0.068)
MSE-F	**9.032**	**15.886**	**22.742**	**25.976**	**23.692**	**36.120**	**38.977**	**27.231**	21.432
(p-value)	(0.000)	(0.001)	(0.000)	(0.002)	(0.014)	(0.012)	(0.018)	(0.035)	(0.058)
Panel B: fixed $e\hat{p}i_t$ vs. C									
MSE_u/MSE_r	0.926	0.879	0.824	0.812	0.800	0.754	0.803	0.921	1.030
ENC-NEW	**16.208**	**25.715**	**32.222**	**36.028**	**27.683**	**33.378**	**32.769**	21.657	11.852
(p-value)	(0.000)	(0.000)	(0.000)	(0.001)	(0.021)	(0.027)	(0.048)	(0.088)	(0.170)
MSE-F	**11.135**	**19.294**	**27.620**	**31.929**	**33.578**	**42.406**	**30.924**	10.417	−3.480
(p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.007)	(0.011)	(0.032)	(0.105)	(0.307)

We use a restricted (benchmark) model of constant returns. The p-values are calculated using a bootstrap based on Kilian (1999). Significant coefficients at the 5% level are highlighted in bold face.

The results indicate that the $e\hat{p}i_t$ forecasting model produces lower MSE than any of the "competitor" model. Moreover, the *MDM* encompassing test indicates that the model using lagged $e\hat{p}i_t$ contains information that provides superior forecasts to those produced by most of the other models. The findings are statistically significant at better than the 2% level in almost every case, regardless of whether the co-integrating parameters are reestimated.[8]

Now, we will examine the predictive power of the log earning–price inflation ratio for long-horizon excess returns. We use the encompassing ENC-NEW test and the equal forecast accuracy MSE-F test presented above. Since these remaining tests have non-standard limiting distributions for overlapping observations that are usually dependent upon unknown nuisance parameters, we follow Clark and McCracken (2004) in using a bootstrap procedure similar to that in Kilian (1999) to estimate asymptotically valid critical values and construct asymptotically valid p-values. We use a restricted model of constant returns.

Table 4 presents out-of-sample statistics of excess returns at horizons ranging from 1 to 24 quarters. The table shows that the unrestricted model (which includes $e\hat{p}i_t$) has smaller MSE than the constant restricted model at horizons less than 6 years. Regardless of whether the co-integrating parameters are reestimated, the ENC-NEW and MSE-F tests reject the null hypothesis that $e\hat{p}i_t$ provides no information about future excess returns at the 5% significance level for horizons of 1 to 16 quarters.

5. Conclusion

The results presented indicate that the earning–price inflation ratio has displayed statistically significant out-of-sample predictive power for excess returns over the post-war period at short and

[8] Except when the log earning–price inflation ratio is recursively reestimated and the competitor models include the consumption–wealth ratio or the dividend–price ratio. However, in these cases, the inverse MDM tests that our variable encompasses the competitor models are not rejected with greater p-value.

intermediate horizons. The one-quarter-ahead non-nested forecast comparison results suggest also that forecasts using $e\hat{p}i_t$ would be consistently superior to forecasts using any other popular forecasting variables.

These results testify to the strong relationship between the earning–price ratio and the inflation level, and then suggest that the decline in inflation since the early 1980s can explain (i) a large part of the run up in equity ratios observed since 1982; and (ii) why econometric tests for structural change provide evidence of a break in the mean financial ratios in the 1990s (Carlson, Pelz and Wohar, 2002; Lettau, Ludvigson and Wachter, 2004).

Our approach is slightly different from previous literature on stock return predictability. In general, the predictability reflects either a time-varying expected risk premium or irrational behavior on the part of market participants. The starting point of our paper is an anomaly, the negative relationship between the price–earning ratio and actual inflation. Instead of exploiting it directly (i.e. when inflation increases then the price–earning ratio decreases and expected returns increase), the predictability comes from the persistence of this anomaly in the long term and the temporary deviations around it.

In that sense, our results are ambivalent concerning the efficient market hypothesis. On one hand, in the long run, real stock prices are attracted by a broad definition of the fundamental value, which include an inflation conditional risk premium. On the other hand, it is difficult to rationalize this feature in a normative model.[9] Our results are rather in line with behavioral finance that has identified a number of cognitive errors to which investors are susceptible. However, the reasons for which inflation makes investors more risk averse remains to be explained. For example, despite the weak empirical evidence on the relationship between inflation and long-run growth, survey questionnaire results presented in Shiller (1997) indicate that nearly 90% of people believe inflation is harmful to economic growth. It could be a piece of the puzzle.

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References

Bakshi, G.S., Chen, Z., 1996. Inflation, asset prices, and the term structure of interest rates in monetary economies. Review of Financial Studies 9, 241–275.
Bruno, M., Easterly, W., 1998. Inflation crises and long-run growth. Journal of Monetary Economics 41, 3–26.
Campbell, J.Y., 1991. A variance decomposition for stock returns. Economic Journal 10 (1), 157–179.
Campbell, J.Y., 2000. Asset pricing at the millennium. Journal of Finance 55, 1515–1567.

[9] In some equilibrium monetary asset pricing models in which money is desired for transaction purposes, real stock returns are negatively correlated with inflation and/or stock prices level negatively correlated with the price level of consumption (e.g. Bakshi and Chen, 1996). But the common long-term trend in the earning–price ratio and actual inflation documented in our paper implies a negative relationship between inflation and required long-run real stock returns.

Campbell, J.Y., 2003. Consumption-based asset pricing, Chapter 13. In: Constantinides, G., Harris, M., Stulz, R. (Eds.), Handbook of the Economics of Finance, vol. IB. North-Holland, Amsterdam, pp. 803–887.

Campbell, J.Y., Shiller, R.J., 1988. The dividend–price ratio and expectations of future dividends and discount factors. Review of Financial Studies 1 (3), 195–228.

Campbell, J.Y., Vuolteenaho, T., 2004. Inflation illusion and stock prices. American Economic Review 94 (2), 19–23.

Carlson, J.B., Pelz, E.A., Wohar, M., 2002. Will the valuation ratios revert to their historical means? Some evidence from breakpoint tests. Journal of Portfolio Management, 22–35 (Summer).

Clark, T.E., McCracken, M.W., 2001. Tests of equal forecast accuracy and encompassing for nested models. Journal of Econometrics 105, 85–110.

Clark, T.E., McCracken, M.W., 2004. Evaluating Long-Horizon Forecasts. Manuscript, University of Missouri at Columbia.

Cochrane, J.H., 1996. A cross-sectional test of an investment-based asset pricing model. Journal of Political Economy 104, 572–621.

Diebold, F.X., Mariano, R.S., 1995. Comparing predictive accuracy. Journal of Business and Economic Statistics 13, 253–263.

Goyal, A., Welch, I., 2003. Predicting the equity premium with dividend ratios. Management Science 49 (5), 639–654.

Harvey, D.I., Leybourne, S.J., Newbold, P., 1998. Tests for forecast encompassing. Journal of Business and Economic Statistics 16 (2), 254–259.

Johansen, S., Juselius, K., 1992. Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for UK. Journal of Econometrics 53, 211–244.

Lamont, O., 1998. Earnings and expected returns. Journal of Finance 53, 1563–1587.

Lettau, M., Ludvigson, S., 2005. Expected returns and expected dividend growth. Journal of Financial Economics 76, 583–626.

Lettau, M., Ludvigson, S., Wachter, J., 2004. The declining risk premium: what role does macroeconomic risk play? Working Paper, New York University.

Modigliani, F., Cohn, R.A., 1979. Inflation, rational valuation and the market. Financial Analysts Journal 35, 24–44.

McCracken, M.W., 2004. Asymptotics for out-of-sample tests of Granger causality, Manuscript, University of Missouri at Columbia.

Phillips, P.C.B., Ouliaris, S., 1990. Asymptotic properties of residual based tests for cointegration. Econometrica 58, 165–193.

Sharpe, S.A., 2002. Reexamining stock valuation and inflation: the implications of analysts' earnings forecasts. Review of Economics and Statistics 84, 632–648.

Shiller, R., 1997. Why do people dislike inflation? In: Romer, C.D., Romer, D. (Eds.), Reducing Inflation: Motivation and Strategy. University of Chicago Press, Chicago.

Stock, J.H., Watson, M.W., 1993. A simple estimator of cointegrating vectors in higher order integrated systems. Econometrica 6, 783–820.