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Reading the Minds of Investors: An Empirical Term Structure Model for Policy Analysis

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Abstract

Building on the recent macro finance literature, this paper develops an empirical term structure model in which investors' judgmental forecasts of macro variables play an important role. The model allows for a limited form of time-variation in the dynamics describing the behavior of short-term interest rates and macro variables. As a result, changes in economic forecasts over time reflect the influence of both economic shocks and perceived changes in economic structure. The latter, in particular, are shown to be important in explaining the evolution of the yield curve over time. An interest rate accounting framework based on the model is applied in parsing changes in long-term interest rates into portions associated with changes in term premiums and changes in expected future short-rates. The changes in expected future short rates are then further decomposed into portions attributable to changes in the expected future paths for inflation, the unemployment rate, and GDP growth and also to a fourth factor interpreted as changes in the "stance of monetary policy." The model results indicate that changes in long-term interest rates, on average, have been about equal parts changes in term premia and changes in expected future short rates. Changes in expected future short rates seem to be driven largely by changes in the stance of monetary policy and in the outlook for inflation while the estimated influence of changes in the outlook for the unemployment rate and GDP growth is more muted.

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[1] The views expressed in this paper are those of the author and do not reflect an official view of the Federal Reserve System. I have benefited from the comments and insights of numerous colleagues including Bill English, Don Kim, Brian Madigan, Bill Nelson, Dick Porter, Vincent Reinhart, and Brian Sack.

Introduction

Every six to eight weeks, a large cadre of dedicated and skilled economists across the Federal Reserve System shift into high gear in readying for their responsibilities over the upcoming FOMC cycle. In the ensuing crush of meetings, model runs, and marathon drafting sessions, economists specializing in seemingly every conceivable facet of the U.S. economy are cross-examined by senior colleagues concerning recent economic and financial market developments and what those developments may hold in store for the future. Among the myriad subjects scrutinized in this way, the recent behavior of interest rates attracts considerable attention. Economists called to the witness stand on this topic might be solicited for their opinions on a range of issues including their sense of the extent to which changes in long-term interest rates over the period since the prior FOMC meeting can be attributed to revisions in inflation expectations, in the outlook for the real economy, and in term premiums. The senior staff inquisitors might also press for some analysis of the likely impact of subtle changes in the wording of FOMC announcements on the behavior of long-term interest rates. From time to time, the impact of "special factors" that have figured prominently in recent market commentary such as the hedging demands of mortgage-backed securities investors or large prospective changes in the supply of Treasury debt might also come under the microscope.

Unfortunately for the staff economists on the firing line, providing plausible answers to such queries based on fully articulated models of the term structure is difficult at best. To be sure, term structure modeling over recent years has witnessed remarkable progress, and the current class of so-called macro-finance models seems to hold great promise as a tool for policy analysts.[2] As yet, however, the formal models that have been developed typically employ assumptions that limit their usefulness in addressing many practical policy issues. For example, interest in policy questions such as those cited above often is most intense following particularly sharp fluctuations in long-term interest rates. These episodes commonly spur a barrage of casual explanations for market developments that implicitly embody an intuition that structural relationships may evolve

[2] The list of papers in this area is multiplying rapidly. Recent important contributions include Ang and Baekert (2004), Ang and Piazzessi (2003), Ang, Piazzessi and Wei (2003), Bernanke, Reinhart and Sack (2004), Hordahl, Tristiani, and Vestin (2003), Kim (2004), Kozicki and Tinsley (2001), and Rudebusch and Wu (2003).

over time and hence that expectations of future macro variables, even at distant horizons, may vary appreciably over intervals as brief as an intermeeting period. However, such intuitive explanations are difficult to reconcile with the fairly simple fixed-coefficient time-series processes assumed to govern the behavior of the short-rate in standard term structure models. In such specifications, changes in expectations about future short-term interest rates and other economic variables can arise only from innovations in the current state of the economy. But absent unit or near-unit roots in the dynamic process for the short-rate, innovations in the current state of the economy typically have little impact on interest rate expectations at distant horizons; current period shocks simply are damped too quickly to have much effect on the longer-term model forecasts. This tension between the quite volatile behavior of interest rates at times and the relative stability of interest rate forecasts from standard econometric models has, of course, been the focus of a sizable body of research over the years beginning with the classic work of Shiller (1979).

Similar in spirit to the work of Fuhrer (1996) and Kozicki and Tinsley (2001), this paper attempts to resolve some of this basic tension by developing a simple empirical term structure model that allows for a limited form of time variation in model parameters. An important byproduct of the model is a set of forecasts for macro variables such as inflation, unemployment, and GDP growth that are consistent with the estimated expected path of short-term interest rates. These forecasts might be viewed as plausible proxies for "market" expectations of these macro variables. The model results indicate that long-term expectations for inflation and other macro variables have evolved over time in a sensible way. Moreover, the model estimates of inflation expectations line up reasonably well with measures of inflation compensation derived from Treasury inflation-indexed securities (TIPS), especially in the last two or three years. We are able to apply the basic model results in a straightforward interest rate accounting exercise that provides an intuitive interpretation of changes in expected future short-rates in terms of changes in expectations for macro variables and changes in the "stance of monetary policy." This accounting exercise suggests that roughly half of the variability in long-term Treasury yields owes to changes in the expected future path of short-term interest rates with the remainder owing to changes in term premiums. Changes in the expected future path of short-term interest rates, in turn, are influenced significantly by changes in inflation expectations and also in the perceived stance of policy. Changes in the prospects for output growth and the unemployment rate also influence interest rate expectations but to a lesser degree.

The remainder of the paper is organized as follows. Section 1 lays out the basic elements of the empirical yield curve model and estimation procedure. Section 2 reviews the results from the baseline model specification. Section 3 describes the interest rate accounting methodology and applies the methodology in reviewing interest rate developments in the summer of 2003. Section 4 offers some concluding comments.

1. An Empirical Yield Curve Model

The empirical model developed below is premised on the notion that market participants' expectations about the future course of interest rates and economic variables are informed both by historical data and by "judgment" about the nature of shocks impinging on the economy and about evolving structural economic relationships. That basic premise has been explored in previous papers including important work by Fuhrer (1996), and Kozicki and Tinsley (2001). But for the most part, previous authors have assumed that perceived structural changes in the economy are manifested entirely in variations in the long-run steady state values of economic variables or perhaps a few other key parameters. While similar in spirit to this previous work, the empirical model developed below is more flexible, allowing for changes in all model coefficients. As a consequence, the evolution in the market's perceptions about the structure of the economy can be reflected in changes in both the estimated steady state values for the model and the model dynamics. To be clear at the outset, the model developed below is not truly a time-varying coefficient model. At any point in time, investors form their expectations of economic variables with a fixed-coefficient model in mind. In what might be described as model drift, however, investors are assumed to update and revise their economic model over time in light of incoming data and judgmental factors.[3]

[3] An excellent example of model drift is discussed in the recent paper by Tetlow and Ironside (2004) that reviews the evolution in the specification of the Board's workhorse macro model – FRB/US – over time.

(1.a) Basic Empirical Model

In contrast to previous work, the empirical model here does not assume that investors employ a particular statistical decision-making framework in updating their beliefs about the structure of the economy. Rather, the model simply assumes that investors rely more heavily upon judgment in developing their forecasts when the historical statistical linkages among the relevant economic variables are weak. Conversely, when historical relationships are very tight, investors are presumed to be less inclined to rely upon judgment in forecasting. The empirical model is then cast essentially as a maximum likelihood or signal extraction problem in which the econometrician, after observing the historical macro data along with the yield curve *for a given day*, attempts to estimate a "plausible" VAR model that might have informed investors' expectations of future short-term interest rates and macro variables on that day. The estimated VAR representation of the economy is a "plausible" model in the sense that: (i) the VAR fits historical data reasonably well and (ii) the VAR forecasts for the short-rate are consistent with the yield curve *on that day*. Note that the emphasis here on matching the yield curve on a particular day is quite different than the approach followed in many so-called factor models of the yield curve, which often seek to identify a small set of underlying, but typically unobserved, stochastic variables that explain the entire history of the yield curve over a long sample. In this regard, the methodology here is closer in spirit to the approach of many investment banks that tune the parameters of their yield curve models each day so as to match the model predictions as closely as possible with the current day's yield curve.

The basic VAR representation for the economy is given by:

$$x_t = c + A(L)x_{t-1} + \varepsilon_t \tag{1.1}$$

where $\varepsilon_t = vector\ of\ shocks\ distributed\ N(0,\Sigma)$

and $x_t = \{r_t, u_t, \pi_t, g_t\}$ is a vector of four economic variables:

$$\begin{aligned}
r_t &= three\text{–}month\ Treasury\ bill\ yield\ (end\ of\ quarter) \\
u_t &= civilian\ unemployment\ rate\ (end\ of\ quarter) \\
\pi_t &= total\ CPI\ \mathrm{inf}lation\ (quarterly, annual\ rate, not\ seasonally\ adjusted) \\
g_t &= real\ GDP\ growth\ (quarterly, annual\ rate)
\end{aligned}$$

These four economic variables are a natural focus for analysis inasmuch as they occupy a central role in a good deal of the macro policy literature in recent years and also lie at the heart of many large macro forecasting models. Indeed, as an alternative to the VAR model specified here, we might instead have posited a simple structural model as in Fuhrer (1996). For our purposes, however, it seemed desirable to remain largely agnostic about the structural model that might have informed investors' expectations of macro variables. In particular, the VAR representation assumed here is general enough to encompass the reduced forms for many simple structural macro models. In the empirical work reported below, x_t, is assumed to follow a first-order VAR process.

The yield curve at a given date *t* is written as a vector of three-month forward rates $\hat{f}_t = \{f_{t+1},...,f_{t+K}\}$ where the subscript *k* indicates the number of quarters ahead; that is, f_{t+k} is the forward three-month Treasury yield *k* quarters ahead. These forward rates ranging from one to eighty quarters ahead are taken from a smoothed Treasury yield curve that is estimated based on the prices for all off-the-run Treasury securities each day.[4] We then assume that the schedule of forward rates can be decomposed into corresponding schedules of expected future spot interest rates, $E_t\{\hat{r}_t\} = \{E_t\{r_{t+1}\},...,E_t\{r_{t+K}\}\}$, forward term premiums, $\hat{p}_t = \{p_{t+1},...,p_{t+K}\}$, and measurement errors, $\hat{\eta} = \{\eta_{t+1},...,\eta_{t+K}\}$ as:

$$\hat{f}_t = E_t\{\hat{r}_t\} + \hat{p}_t + \hat{\eta}_t \qquad (1.2)$$

To close the model, the forward term premium schedule for date *t* is written as a simple polynomial in the variable *k* representing the number of quarters ahead. That is, the k[th] element of $\hat{p}_t$ is written as:

$$p_{t+k} = \alpha_1 k + \alpha_2 k^2 + \alpha_3 k^3 + \alpha_4 k^4 \qquad (1.3)$$

(Note that the polynomial specification for the term premium does not include a constant term, thus appropriately restricting the term premium in the current period (k=0) to be equal to zero.) The polynomial specification (1.3) places far less structure on the term premium specification than is common in much of the recent term structure literature, which has devoted considerable attention to structural modeling of the time-series

[4] The estimation employs a standard parametric specification for the forward rate curve, Svennsson (1994).

behavior and economic determinants of risk premiums, Dai and Singleton (2003). A common assumption in the no-arbitrage term structure literature holds that the stochastic discount factor employed in valuing future cash flows can be written as a function of the innovations in the underlying state variables. On any given day, however, the term premium schedule implied by that structural approach still reduces to a (very complicated) function of time to maturity, *k,* and other parameters of the model, Ang and Piazzessi (2003). In effect, the approach here approximates that function by a simple low order polynomial. Of course, the absence of an explicit structural specification for the risk premium in the model entails some significant costs. In particular, the estimated model does not generate estimates of the market prices of various risk factors and hence we sacrifice the ability to use an estimated pricing kernel in valuing other financial assets. However, the less structural approach to estimating term premiums in this paper seems defensible given our objectives. Unlike most of the extant macro-finance literature, here we are primarily interested in identifying changes over time in the market's perceptions of the process driving the macro variables. Imposing strong and possibly mis-specified no-arbitrage restrictions in modeling the term premium would likely tend to induce greater and possibly spurious time variation in the estimated macro VAR coefficients.

(1.b) Likelihood Function

Assuming that the VAR shocks and the yield curve measurement errors are uncorrelated, the joint log likelihood function implied by the VAR model (1.1) and the yield curve specification (1.2), is given by:

$$-2\log(L) = T\log(|\Sigma|) \quad + \quad K\log(\sigma_\eta^2) \quad + \quad \sum_{t=1}^{T}\varepsilon_t'\Sigma^{-1}\varepsilon_t \quad + \quad \sum_{k=1}^{K}(\eta_{t+k}/\sigma_\eta)^2 \tag{1.4}$$

The third term in this expression just represents the fit of the VAR over historical data. The last term in the expression represents the fit of the VAR forecasts with the *current day's* yield curve. The macro data employed in this estimation are quarterly, but yield curves, of course, are available daily. To estimate the model on days beyond the most recent quarterly observation of macro data but prior to the release of current quarter macro data, it seems reasonable to amend equation (1.4) to allow for shocks in the "contemporaneous" period T+1. The notion is simply that market participants respond to news and information every day, and that information is reflected in the shape of the yield curve. In part, changes in the yield curve may reflect updates in beliefs about the structural coefficients of the model. But daily fluctuations in the yield curve no doubt also reflect updated estimates of the current state of the economy attributable to various shocks that have been observed to date during the quarter. To allow for this effect, we include the additional contemporaneous term in the likelihood function:

$$-2\log(L) = T\log(|\Sigma|) + K\log(\sigma_\eta^2) + \left(\sum_{t=1}^{T}\varepsilon_t'\Sigma^{-1}\varepsilon_t + \varepsilon_{T+1}'\Sigma^{-1}\varepsilon_{T+1}\right) + \sum_{k=1}^{K}(\eta_{t+k}/\sigma_\eta)^2 \tag{1.5}$$

Equation (1.5) is the core likelihood function specification. In the estimation, the shocks in the contemporaneous period T+1 are treated simply as additional model parameters to be estimated. [5]

(1.c) Constraints

There are no guarantees that the VAR obtained minimizing the likelihood function (1.5) will be a "plausible" representation of the economy from a macroeconomist's perspective. For example, the estimated VAR might imply explosive dynamics or forecasts for the short-rate that are negative over some range. In addition, in the attempt to match the current day's yield curve very closely, minimization of (1.5) could result in VAR parameter estimates that are quite at odds with historical data. To ensure that the VAR obtained in estimating (1.5) is plausible in this broader sense, we append four penalty terms to the core objective function in the estimation. The first penalty function ensures that the estimated VAR is dynamically stable. To impose the dynamic stability condition, the stability term penalizes VAR specifications in which the maximum of the absolute values of the eigenvalues for the VAR transition matrix A exceeds 0.90. (For complex eigenvalues, the penalty is based on the magnitude of the complex absolute value of the eigenvalue). Setting a maximum for the absolute values of

[5] Note that the variance-covariance matrix for the contemporaneous error vector is assumed identical to that applicable for the historical errors. There is an argument that the contemporaneous variance-covariance matrix should be scaled in some way to reflect the fact that, in most cases, a full quarter will not have passed since the end of the previous quarter.

the eigenvalues at this level ensures that any shock to the VAR will largely damp out after about ten years. The penalty function thus takes the form:

$$\begin{aligned}&\textit{Stability Penalty} = \theta \max(0, \max(\operatorname{abs}(\omega)) - .9)\\&\textit{where} \quad \omega = \textit{vector of eigenvalues of VAR transition matrix } A\end{aligned}$$

Of course, there is nothing in the estimation that prevents the estimated VAR coefficients from taking on values that imply a more rapid dynamic convergence if such a specification improves the model fit. With the scale factor, θ, in the penalty function set sufficiently high, minimization of the objective function will result in estimated eigenvalues with absolute values that are, at most, just a touch above 0.9, implying that the value of the penalty term itself will be very close zero at the optimum parameters.

We also would like to avoid models that imply forecasts for the short-term rate that violate the zero bound on nominal interest rates. To this end, we add a second penalty term to the objective function that punishes any VAR specifications that produce forecasts of the short-rate that are negative at any point over the forecast horizon. Thus the zero bound penalty function takes the form:

$$\textit{Zero Bound Penalty} = -\theta \sum_k \min(0, E_t\{r_{t+k}\})$$

Note that the form of this penalty (and all the other penalty terms) implies that no costs are incurred until the bound is actually exceeded, so the presence of the penalty term in the optimization does not affect the value of "interior" solutions.

Finally, we would like to ensure that the estimated VAR is reasonably consistent with historical data. Two additional penalty terms are added to the objective function for this purpose. The first term penalizes "large" deviations in the estimated VAR coefficient vector from the corresponding least squares estimate of the coefficient vector. The quadratic form $Q_\beta = (\beta - \beta_{ls})\bar{V}^{-1}(\beta - \beta_{ls})$ provides a natural metric of the degree to which the coefficient estimates are at odds with historical data. In this expression, the vector β is the "restricted" VAR coefficient vector estimated to match the yield curve and β_{ls} is the unrestricted least squares estimate of the VAR coefficient vector. The matrix $\bar{V}$ is the unrestricted least squares estimate of the variance-covariance matrix for the VAR coefficient estimates. With a total of twenty estimated VAR coefficients (four equations with five estimated parameters for each), the statistic Q_β is distributed $\chi^2(20)$ and the penalty function for the coefficient vector is thus specified as:

$$\textit{Coefficient Penalty} = -\theta \min(0, Q_{10} - Q_\beta)$$
$$\textit{where } Q_{10} = 10 \textit{ percent critical value for } \chi(20) \textit{ distribution}$$

The intuition is simply that the coefficient penalty term punishes any VAR coefficients that would be rejected by the historical data at the ten percent confidence level. Again, with the scale factor, θ, set at a high value, the penalty term itself takes on a value very close to zero at the estimated parameter values.

We would also like to ensure that the estimation results in a set of contemporaneous macro shocks that is consistent with the distribution of historical shocks. For this purpose, we add a similar penalty term for the contemporaneous errors. The quadratic form $Q_\varepsilon = \varepsilon'_{T+1}\bar{\Sigma}^{-1}\varepsilon_{T+1}$ is distributed $\chi^2(4)$ and large values of this statistic would indicate shocks that are at odds with the variance-covariance patterns evident in the historical unrestricted VAR residuals. In this expression, the matrix $\bar{\Sigma}$ is the unrestricted least squares estimate of the variance-covariance for the VAR residuals. The analogous shock penalty term is then given by:

$$\textit{Contemporaneous Shock Penalty} = -\theta \min(0, Q_{10} - Q_\varepsilon)$$
$$\textit{where } Q_{10} = 10 \textit{ percent critical value for } \chi(4) \textit{ distribution}$$

Thus the full estimation chooses the VAR coefficients c and A, term premium coefficients, $\alpha_1,..,\alpha_4$, and contemporaneous shocks, ε_{T+1}, to minimize the expression:

$$\begin{aligned}
&-2\log(L) = \\
&T\log(|\Sigma|) \quad + \quad K\log(\sigma_\eta^2) \quad + \quad \left(\sum_{t=1}^{T}\varepsilon_t'\Sigma^{-1}\varepsilon_t + \varepsilon'_{T+1}\Sigma^{-1}\varepsilon_{T+1}\right) \quad + \quad \sum_{k=1}^{K}(\eta_{t+k}/\sigma_\eta)^2 \\
&+ \quad \textit{Stability Penalty} \quad + \quad \textit{Zero Bound Penalty} \quad + \\
&\quad \textit{Coefficient Penalty} \quad + \quad \textit{Contemporaneous Shock Penalty}
\end{aligned} \tag{1.6}$$

As a final wrinkle in the objective function, the dynamic stability requirement for the estimated VAR ensures that the expected path for the short-rate will be essentially flat at the long-run level after about ten years. This suggests that we should expect the model to fit the forward rate schedule especially closely at horizons beyond ten years. To capture this effect, we split the yield curve residuals into groups—those for the first ten years and those for the second ten years. We assume that the standard error of the yield curve residuals for the first ten years, σ'_η, is 0.05 percent (five basis points) while the standard error of yield curve residuals beyond ten years, σ''_η, is set at 0.005 percent (one-half basis point). This entails a minor additional modification in the objective function as:

$$-2\log(L) =$$

$$T\log(|\Sigma|) \quad + \quad K\log(\sigma_\eta^2) \quad + \quad \left(\sum_{t=1}^{T}\varepsilon_t'\Sigma^{-1}\varepsilon_t + \varepsilon_{T+1}'\Sigma^{-1}\varepsilon_{T+1}\right) \quad + \quad \sum_{k=1}^{40}(\eta_{t+k}/\sigma'_\eta)^2 + \quad \sum_{k=41}^{80}(\eta_{t+k}/\sigma''_\eta)^2$$

$$+ \quad \textit{Stability Penalty} \quad + \quad \textit{Zero Bound Penalty} \quad + \quad \textit{Coefficient Penalty} \quad + \quad \textit{Contemporaneous Shock Penalty} \tag{1.7}$$

(1.d) Data Smoothing and Estimation Periods

Apart from the model specification, the treatment of noise in underlying macro data and the choice of estimation periods are also important aspects of the model. We assume that the raw macro data are noisy and that investors attempt to strip out this noise by applying a Kalman filter before making their forecasts. The smoothing algorithm here assumes that there is a set of "underlying" macro state variables driven by the VAR specification (1.1) assumed above:

$$x_t = c + A(L)x_{t-1} + \varepsilon_t$$

However, the observed macro variables are noisy readings on these underlying state variables:

$$\tilde{x}_t = x_t + u_t$$

where u_t is a vector of measurement errors distributed $N(0,\Lambda)$. The measurement errors are assumed to be uncorrelated across variables so the off diagonal elements of the variance-covariance matrix, Λ, are all zeros. To smooth the data, we assume that the standard deviations of the fundamental shocks to the state variables, ε_t, are equal to one-half of the standard deviations of the corresponding measurement error—that is, $diag(\Sigma) = 0.5 \cdot diag(\Lambda)$. With this assumption, the Kalman filtering algorithm calculates maximum likelihood estimates of the underlying state variables. These smoothed data then are taken as the raw inputs in the estimation.

With the filtered macro data in hand, another important modeling element concerns the choice of estimation period. In some respects, it seems natural to assume that in forming their expectations, individuals might tend to weight recent observations more heavily than those in the distant past. This line of reasoning might argue for some type of weighting scheme in the estimation that would specify the standard deviation of the VAR errors for each observation as an increasing function of the period of time between that historical observation and the current date. After some experimentation, the discussion below adopted the simpler approach of estimating the model over moving twenty-year estimation periods. That is, for each yield curve, the estimation fits the macro VAR using historical data over the previous twenty years.

2. Basic Model Results

In the baseline model, we assume a simple first-order VAR specification for the macro variables. Including additional lags in the VAR is feasible, but comes at a substantial cost in terms of computational burden. This section presents the basic results obtained by estimating the model to match, in sequence, each yield curve prevailing on the dates of all FOMC meetings since December of 1987.

(2.a) Data Smoothing

As a first order of business, it seems useful to examine the smoothed macro data that are the key inputs for the model. Exhibit 1 provides a sense of the difference between the raw data and the filtered data using the technique described above. In general, this application of the Kalman filter seems to be effective in knocking out much of the very high frequency volatility in each series while preserving the broader time-series profile of each variable. As one might expect, the estimated "underlying" variables for the short-rate and for the unemployment rate are very close to the "observed" values for these variables. In contrast, total CPI inflation and GDP growth are very volatile at a quarterly frequency and the Kalman filter determines that a sizable portion of the quarter-to-quarter fluctuations in these variables is attributable to "noise" that would not be expected to persist.

(2.b) Market-Consistent Forecasts of the Short-Rate and Other Economic Variables

Exhibit 2 displays a sample of the basic yield curve decomposition. The upper left panel of the chart displays the forward rate curves on the dates of the June FOMC meetings in 2003 and 2004. For these two dates, the VAR estimated by minimizing (1.7) generates corresponding forecast schedules for the short-rate shown in the upper right panel. The difference between the forward rate and the expected short-rate curves represents the term premium schedule (lower left panel) for each date—up to an allowance for a small residual (lower right panel). The term premium schedules tend to rise monotonically for the dates shown. The model generally fits the yield curve very well, although the residuals point to some difficulty in fitting the curve in the very near-term. Table 1 below reports the standard deviation of the model residuals for zero coupon yields ranging from one to ten years over the full sample period from December 1987 through August 2004. Of course, by construction, the model fits observed yields over time considerably more closely than no-arbitrage term structure models that employ only observable macro factors and a fixed-coefficient time-series process for the macro dynamics.[6]

Table 1: Standard Deviation of Model Residuals at Selected Maturities

Zero Coupon Yields	Standard Deviation (Basis Points)
One-Year	6.7
Two-Year	4.8
Three-Year	2.5
Four-Year	3.0
Five-Year	1.7
Seven-Year	1.7
Ten-Year	0.8

Several features of the estimation results are apparent in the panel for the expected short-rate in exhibit 2. First, expectations of the near-term path for the short-rate seem to vary quite a bit over time. In June 2003, for example, market participants

[6] Bernanke, Reinhart, and Sack (2004) employ such a model and, as one would expect for models of this type, the authors report residual standard errors much larger than those noted in Table 1.

were reportedly quite concerned about the economic outlook. Measures of core price inflation had dropped to unusually low levels and the risk of further substantial disinflation seemed very real given the apparent slack in labor markets and sluggish output growth. Indeed, the Chairman's testimony before Congress in April of that year along with the speeches of other Federal Reserve officials had sounded warnings about the dangers of "unwelcome further disinflation" and even broached the possibility, albeit quite remote, that "unconventional" monetary policy might become necessary to avoid a Japan-like deflationary scenario. Not surprisingly then, the expected path of the short-rate on this date exhibits an expected easing of policy in the very near term followed by a very gradual anticipated tightening of policy. A year later in June of 2004, the economy seemed to be on a much stronger footing. As shown in the top right panel, the model suggests that the expected path for the short-rate shifted up substantially over this period with investors anticipating a significant degree of tightening over the next two years. The model estimates of the term premium schedules for these two dates—shown in the bottom left panel—are broadly similar and suggest that the term premium in each case rises fairly steadily to a level of about 2½ percentage points at the ten-year (40 quarter) horizon. At intermediate maturities, the term premium schedule for June of 2004 lies noticeably above that in June of 2003, perhaps indicating the waning of safe-haven demands for Treasury securities as concerns about especially adverse economic scenarios faded over time.

Exhibit 3 displays the time series for this forward rate decomposition at selected horizons. As one would expect, most of the variation in near-term forward rates is attributed to changes in the expected short-rate. At longer horizons, a larger portion of the variability in forward rates is associated with movements in the term premium. Nonetheless, the time-series variability in the ten-year ahead expected future short rate is also significant. The model suggests that term premiums at the selected horizons have been positive, on average. However, there are some periods—particularly in 1998—during which the estimated term premium schedule takes on small negative values, even at distant horizons. One interpretation of this result is that during periods of severe financial distress like that precipitated by the Russian debt default and LTCM debacle, investors seeking a safe harbor from market turmoil may value risk-free and highly liquid assets such as Treasury securities at a substantial premium, resulting in elevated bond prices and depressed yields. Absent a variable in the VAR that adequately captures such special safe haven demands, the unusually elevated bond prices during such scenarios would show up as very low or even negative term premiums.

The estimated yield-curve-consistent VAR for each date can be used to generate forecasts for the other variables in the model as well. As shown in the middle right panel of exhibit 4, these forecasts suggest that the expected trajectory of inflation has been marked up over the last year. Much of the upward revision has been concentrated in relatively near-term horizons. This finding might be viewed as consistent with the view that the risks of substantial disinflation that were the focus of so much attention in the months leading up to the June 2003 FOMC meeting are now much attenuated in light of the generally improved economic outlook and the uptick in core price inflation observed over the first half of this year. It is interesting to note that—market speculation about possible deflation scenarios notwithstanding—the model results suggest that investors anticipated a fairly flat trajectory for inflation in June 2003. However, as noted in the bottom left panel, perceptions of an unusually accommodative monetary policy seem to have been an important factor stabilizing inflation expectations at this time. For example, the model suggests that in the summer of 2003, market participants believed the FOMC would maintain a negative real federal funds rate for almost five years.

The sense that the market outlook for the economy has become more sanguine over the last year is evident in the charts for the expected paths of the unemployment rate and GDP growth. These forecasts suggest that as of June 2004, investors had marked up their estimates for GDP growth relative to last year and looked for the unemployment rate to remain close to 5½ percent for the foreseeable future—a half percentage point or more below expectations for the unemployment rate in June of 2003.

The qualitative features of these results tend to support the research findings of other authors—Gurkaynak, Sack, Swanson (2003) among others—that long-term inflation expectations can respond significantly to economic developments. However, the estimates of long-term inflation expectations are not particularly volatile in the short-run and thus do not seem to accord well with historical interpretations that appeal to abrupt changes in inflation expectations as the dominant factor explaining high frequency movements in long-term interest rates, Goodfriend (1994), Ang and Bekaert (2003).

Exhibit 5 compares the model-implied inflation expectations with inflation compensation implied by yields on Treasury inflation-protected securities (TIPS).[7] TIPS inflation compensation, of course, embeds inflation expectations as well as a premium for inflation risk. Consistent with the discussion in Sack and Elsasser (2003), and Kim (2004), the model suggests that TIPS inflation compensation ran below inflation expectations in the early years of the program, implying a negative inflation risk premium. However, as the market has matured over more recent years, TIPS inflation compensation has moved up and now nearly matches the level of model-implied expected inflation rates. Apart from this secular trend, the chart also suggests some positive correlation between changes in TIPS inflation compensation and changes in the model-implied expected inflation rates. As shown in the scatter plot in the right panel, the red fitted regression line indicates that a 100 basis point change in five-year TIPS inflation compensation would be associated with about a 30 basis point change in model-implied five-year inflation expectations. Thus the model suggests that much of the very short-run variation in TIPS inflation compensation seems to be attributable to changes in inflation risk premiums rather than inflation expectations.

Exhibit 6 presents the time-series of the forecasts for the short-rate, inflation, and real interest rates (defined as the expected short-rate less the expected inflation rate). At each horizon, there is substantial time-series variation in the expected future short-rate. The model results suggest that some of this variation in expected future short rates is associated with movements in inflation expectations in the same direction. But the model results also point to considerable time-series variation in expected real short-term interest rates at all horizons. The results also indicate that changes in expected real short-term interest rates are generally positively correlated with movements in expected future nominal short-rates and inflation. One might expect to observe such correlations, for example, if investors viewed policy as roughly following a Taylor-type feedback rule.

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[7] To a first approximation, inflation compensation is often calculated essentially as the spread between a nominal Treasury security and a comparable-maturity inflation indexed security. More sophisticated calculations take account of the backloaded nature of cash flows from TIPS securities relative to standard nominal securities.

Exhibit 7 displays the time-series for the forecasts of the macro variables at selected horizons. As one would expect, the near-term forecasts for all the variables are considerably more variable than distant-horizon forecasts. In most cases, the near-term forecasts tend to fluctuate around the level of the long-term forecast. For the unemployment rate, however, the model estimates suggest that investors were nearly always anticipating the unemployment rate to rise over time—that is, the near-term unemployment rate forecast tends to run below the long-term forecast. In part, this seems to reflect changing perceptions of the interest sensitivity of the unemployment rate, particularly in the early 1990s. At that time, the model suggests that investors were anticipating that the sharp decline in short-term interest rates would, as in the past, be followed shortly by a surge in output growth and a fairly steep drop in the unemployment rate. In fact, however, the unemployment rate fell rather gradually, in part perhaps reflecting the difficulties in the financial sector and other economic “headwinds.” The model interprets these years as a period in which investors persistently underestimated the forces hindering the recovery in labor markets.

(2.c) VAR Asymptotes

Exhibit 8 displays the time variation in the estimated VAR asymptotes for the four model variables. The VAR asymptotes are of considerable inherent interest as indicators of the steady state tendencies of the economy, which are factors that often figure prominently in economic forecasting and in policy deliberations. For example, the asymptote for GDP growth might be viewed as a measure of the markets’ view of long-run potential output growth for the economy. Similarly, the asymptote for the inflation rate might be taken as a measure of investors’ sense of the central bank’s “inflation target.” The asymptotes for the nominal short-rate together with the estimated long-run inflation rate yield a measure of the long-run real short-rate, which might be regarded as a measure of investors’ assessment of the long-run “equilibrium” real interest rate. And finally, the asymptote for the unemployment rate might be viewed as investors’ sense of the long-run NAIRU.

In general, the model results indicate that the long-run GDP growth rate has edged up over time from around 3 percent to about 3½ in recent years. That result seems consonant with much analysis over recent years identifying a substantial pickup in the trend growth of labor productivity. The long-run inflation rate has edged lower from a range of 4½ to 5½ percent in the late 1980s and early 1990s to about 3 percent now with much of that downshift occurring in the last few years. Investors' estimate of the long-run NAIRU has likewise edged lower over time with the most recent reading in the vicinity of 5½ percent. The model estimates of the long-run nominal short-rate have moved substantially lower over time as well to a level of about 3½ percent in June 2004, evidently partly reflecting investors' sense of the improvement in long-run inflation prospects. The expected long-run real interest rate has apparently moved lower as well. For much of the 1990s, the model suggests that investors believed the long-run level for this measure of the real interest rate fluctuated in a range between 1½ and 2 percent. With some estimates suggesting that total CPI inflation might overstate "true" inflation by perhaps a percentage point, a steady state real interest rate in the 1½ to 2 percent range would then imply that the "true" equilibrium real interest rate over much of the 1990s ranged between 2½ and 3 percent. In recent years, however, the model asymptote for the short-term real rate has dropped substantially to a level of only about ¾ percent. Assuming that the market's perception of the CPI inflation bias has not changed much of late, the model asymptote for the short-term real rate of ¾ percent would suggest that investors now put the equilibrium short-term real rate at only about 1¾ percent—a level that is rather low by historical standards but within the range of many other recent estimates of equilibrium real interest rates.

(2.d) VAR Dynamics

The preceding discussion of the model asymptotes focused essentially on the time-variation in the estimated constant terms in each equation of the yield-curve-consistent VAR. However, the estimates of all the model parameters are allowed to change over time, so one might expect to observe some time variation in the implicit VAR dynamics. Impulse response patterns are a convenient means of characterizing this evolution in the estimated model dynamics. Exhibit 9 depicts the response of each of the model variables to a 1 percentage point "own equation" shock. For example, the upper left panel displays the impulse response pattern of the short-rate to a 1 percentage point positive shock to the short-rate equation; the upper right panel displays the impulse response pattern of the unemployment rate to a 1 percentage point positive shock to the unemployment equation and so on. These charts suggest that investors' believed the shocks affecting the short-rate in June 2003 were relatively persistent. In part, the perception of heightened persistence of the short-rate process in 2003 could reflect the efforts of the FOMC in conveying a sense that the stance of monetary policy was likely to remain accommodative for some time. For example, the June 2003 FOMC announcement included a statement indicating that the possibility of an "unwelcome substantial fall in inflation" would likely be its predominant concern for the foreseeable future.[8]

Exhibit 10 displays the full impulse response pattern over time. The shocks here have not been orthogonalized nor has there been any attempt to identify structural shocks. Nonetheless, the impulse response patterns appear fairly reasonable. For example, moving across the top row, the unemployment rate typically moves up following a positive interest rate shock while inflation and GDP growth move lower. And moving down the first column, the short rate tends to rise in response to inflation and growth shocks and to fall initially in response to a positive unemployment shock. Although the dynamic stability penalty in the estimation ensures that the effects of all shocks dissipate almost completely after about ten years, market participants' views about the speed of convergence for the economy do seem to vary significantly over time.

(2.e) VAR Forecast Variance

Macroeconomists and policymakers are often interested in the markets' perceptions of uncertainty about the economic outlook. The VAR framework employed here is convenient for this purpose as well. The forecast variance for the model variables is given by:

$$\Psi(k) = E_t(x_{t+k} - E_t\{x_{t+k}\})(x_{t+k} - E_t\{x_{t+k}\})' \tag{2.1}$$

This expression is a complicated function of the estimated VAR parameters and the error-term variance-covariance matrix. In a standard fixed-coefficient VAR, the forecast variance function is the same at all points over the estimation period. In the model above, however, the estimates of the model's parameters vary considerably over time. As these VAR coefficients change, the implied forecast variance function changes as well. In

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[8] See Kohn and Sack (2003), Gurkaynak, Swanson, and Sack (2004), and Bernanke, Reinhart, and Sack (2004) for discussions of the important role of FOMC communications in this episode.

particular, the variation in the forecast variance function is directly related to the variation in the impulse response patterns discussed above. When the model coefficients imply that shocks are quickly damped, the forecast variance function will tend to reach an asymptote quickly at a comparatively low level. Conversely, when shocks are viewed as quite persistent, the forecast variance function reaches a higher asymptote and rises along a flatter trajectory toward this higher level.

Exhibit 11 displays the standard deviations of forecast errors for the short-rate, unemployment rate, the inflation rate, and GDP growth on the dates of the June FOMC meetings in the past two years. In general, these results suggest that investors' sense of uncertainty about the macroeconomic outlook has not changed greatly over the past year. As shown in the bottom left panel, uncertainty about inflation seems to be something of an exception with the model results pointing to a modest upturn in inflation uncertainty over the past year.

Exhibit 12 displays the time-series for these forecast error standard deviations at selected horizons. A prominent feature of these charts is the substantial decline in forecast uncertainty over time recorded for all four variables over the last two decades. Of course, the apparent decline in the volatility of macroeconomic variables in recent decades has been well documented, Kahn, McConnell, Quiros-Perez (2002), Stock and Watson (2002). Curiously, though, the model estimates of term premiums in exhibit 3 do not show any corresponding secular decline. The absence of any notable diminution in investors' required compensation for interest rate risk suggest that the market prices of macro risk may have risen over time or perhaps that additional factors apart from the macro variables considered here may also be important determinants of interest rate risk premiums.

(2.f) Estimated Shocks

As described above, the estimation procedure infers shocks that have hit the economy in the period since the last quarterly observations of macro data. Exhibit 13 plots the estimates of these estimated shocks over time. A prominent feature of this chart is the strong positive correlation over time in the estimated shocks for the short-rate, the inflation rate and GDP growth. Conversely, there is a strong negative correlation over time in the shocks for these three variables with the estimated shocks for the unemployment rate. These correlations arise naturally in minimizing the objective function (1.7). The model chooses the shock for the short-rate and other variables to match the forward rate in the contemporaneous period and in subsequent periods. But the shocks to the unemployment rate, inflation rate, and GDP growth do not affect the forecast for the short-rate in the contemporaneous period. As a result, the interest rate shock is chosen largely to minimize the yield curve error in the contemporaneous period. As an informative special case, one can readily compute the expected value of the shocks for the other macro variables conditional on this "observed" interest rate shock and the shock covariance matrix Σ. In this case, the estimated shocks to the macro variables would be perfectly correlated with the estimated interest rate shock. Of course, the full model estimation of the macro shocks is more complicated than this simple conditional expectation calculation. But in many cases, the estimated shocks for the unemployment rate, inflation and GDP growth from the full estimation are very similar to the simple conditional expected values of the shocks.

In scanning the history of the estimated shocks, two episodes stand out as especially noteworthy. In 1994, the Federal Reserve tightened policy sharply following an extended period during which the target funds rate was maintained at 3 percent. The model interprets a portion of this tightening as a sequence of large positive interest rate shocks. In effect, one might view such shocks as positive deviations from a Taylor rule of sorts—e.g. interest rate movements that are not well explained by observed values of macro variables. Conversely, the FOMC eased policy very aggressively in early 2001 to address the incipient weakening in economic activity. The model attributes some portion of this sharp drop in rates to negative interest rate shocks. This finding again suggests that the easing in 2001 was more aggressive than what investors might have expected based upon their perceptions at the time of the policy feedback rule describing FOMC behavior.

(2.g) Model Diagnostics

Exhibit 14 provides some perspective on the overall fit of the model. As described above, a simple chi-squared statistic provides a useful gauge of the extent to which the estimated VAR coefficients depart from those obtained from unconstrained least squares regression estimates. As shown in the upper left panel, the chi-squared statistic for the regression coefficients never approaches the ten percent critical level. However, the somewhat higher level of this statistic in the last few years suggests that investors' implicit view of the structure of the economy has departed more noticeably from that suggested by the unconstrained least squares VAR estimate. The upper right panel displays the analogous chi-squared statistic for the estimated contemporaneous shocks. The value of the statistic suggests that investors might have viewed the economy as subject to rather large shocks over time. That said, the chi-squared statistic reaches the ten percent critical level on only a few occasions.

3. An Application: Interest Rate Accounting

Having now surveyed at some length the core features of the model results, we return to a theme raised at the outset—the challenges confronting Federal Reserve economists that have been called upon to make sense of interest rate developments over the intermeeting period. As noted above, the model results offer a rich body of information that can potentially be useful in interpreting the historical behavior of the yield curve including estimates of term premiums along with "plausible" yield-curve-consistent measures of the expected future paths for the short rate, inflation, the unemployment rate and GDP growth. To push the policy analysis a step farther, we consider a simple tool that might be described as interest rate accounting. Given the estimated expected future paths of macro variables at various dates and the corresponding estimated VAR equations, we can ask what portion of any observed change in interest rates owes directly, in an accounting sense, to changes in the expected macroeconomic outlook for the economy. The starting point for the interest rate accounting exercise is the estimated interest rate equation from the VAR given by:

$$\hat{r}_t = \hat{a}_0 + \hat{a}_1 r_{t-1} + \hat{a}_2 u_{t-1} + \hat{a}_3 \pi_{t-1} + \hat{a}_4 g_{t-1} + \hat{\varepsilon}_t \tag{3.1}$$

The expected short rate at any future date t+k can then be written as a function of the expected future paths of the macro variables as:

$$\hat{r}_{t+k} = (\hat{a}_0 / \hat{a}_1)\cdot(1-\hat{a}_1^{k+1}) + (\hat{a}_1)^k \hat{r}_t + \sum_{s=1}^{k} (\hat{a}_1)^{k-s} \cdot (\hat{a}_2 \hat{u}_{t+s-1} + \hat{a}_3 \hat{\pi}_{t+s-1} + \hat{a}_4 \hat{g}_{t+s-1}) \tag{3.2}$$

The expected future path of short rates estimated at some later date t' is calculated in analogous fashion. To measure the arithmetic contribution of the revision in the expected paths for macro variables to the revision to the expected future path of short-rates, we hold the coefficients in equation 3.1 constant at the estimates from date *t,* and compute the macro revision contributions as:

$$period\ k\ unemployment\ rate\ revision\ contribution = \sum_{s=1}^{k} (\hat{a}_1)^{k-s} \cdot \hat{a}_2 \cdot (\hat{u}_{t'+s-1} - \hat{u}_{t+s-1})$$

$$period\ k\ inflation\ revision\ contribution = \sum_{s=1}^{k} (\hat{a}_1)^{k-s} \cdot \hat{a}_3 \cdot (\hat{\pi}_{t'+s-1} - \hat{\pi}_{t+s-1})$$

$$period\ k\ GDP\ growth\ revision\ contribution = \sum_{s=1}^{k} (\hat{a}_1)^{k-s} \cdot \hat{a}_4 \cdot (\hat{g}_{t'+s-1} - \hat{g}_{t+s-1})$$

The impact of changes in the initial short rate and changes in the regression coefficients between dates *t* and *t'*—which we label the "policy stance" revision contribution—is computed as a residual by subtracting the sum of the individual macro revision contributions described above from the estimated total revision in the expected future path of the short rate. That is,

$$policy\ stance\ revision\ contribution = \{\hat{r}_{t'+k} - \hat{r}_{t+k}\} - \{period\ k\ macro\ revisions\ contributions\}$$

We call this the policy stance revision contribution because it captures the impact of changes in the current period level of the short rate coupled with the effects of changes in the coefficients for the interest rate equation in the VAR. For *given* expected paths for inflation, the unemployment rate and output growth, the changes in the interest rate equation coefficients measure the extent to which investors anticipate a higher or lower path for the short-rate. Thus, the concept of the revision in "policy stance" employed here might be viewed as capturing the exogenous tightening or easing implicit in the steeper or flatter trajectory for the expected short-rate path.

The upshot of this analysis is that the revisions to the expected future path of the short rate can be decomposed into portions corresponding to the arithmetic contributions of changes in the outlook for macro variables and changes in the stance of policy. Adding the revisions in the term premium schedule between the two dates then provides a full arithmetic decomposition of the corresponding forward rate revisions.

Using this methodology, exhibit 15 displays the interest rate accounting decomposition for the change in the ten-year zero coupon Treasury yield over the June 2003 intermeeting period (e.g. the period between the May and June FOMC meetings) and for the August 2003 intermeeting period (the period between the June and August FOMC meetings). As noted above, market concerns about potential deflationary scenarios helped to push long-term yields down substantially over the June intermeeting period. Thereafter, long-term rates rebounded sharply on more positive economic news and, reportedly, some disappointment among investors that previously had attached at least some probability to scenarios in which the Federal Reserve would resort to "unconventional" monetary policy operations in an effort to drive down long-term yields. The arithmetic decomposition for the June intermeeting period suggests that a change in the inflation outlook and in the perceived stance of policy did, indeed, push yields lower. A decline in the term premiums was also a major factor behind the drop in yields. Of course, the term premium estimates in this model encompass any factors influencing yields that are not captured directly by the variables included in the VAR. As such, the sizable negative contribution of term premiums over the June intermeeting period could be viewed as lending credence to ubiquitous reports in the financial press at that time suggesting that the initial drop in long-term yields was amplified by the hedging activities of mortgage-backed securities investors as they flocked to purchase long-term Treasuries in an attempt to grapple with rising prepayment speeds and the attendant drop in their portfolio durations.

The arithmetic revision contribution for the August 2003 FOMC meeting is even more striking. Incoming economic data over July and early August generally came in on the high side of expectations and, as shown by the orange bar, the GDP growth revision contribution was positive. But the model suggests that the lion's share of the sharp jump in long-term rates over this interval owed to a very sharp rise in term premiums. It is possible that the rise in term premiums might have been occasioned by an increase in interest rate volatility—the model estimates of the interest rate volatility curve did edge up from the June to the August meeting. But this measure of volatility also climbed a bit from May to June when the estimated term premiums declined. Evidently, fluctuations in the market price of interest rate risk, again perhaps attributable to the portfolio balancing strategies of mortgage investors, were a key factor driving the estimated wide swing in term premiums between the May and August meetings.

The revision decompositions over these two intermeeting periods are atypical in some respects. Table 2 below, for example, computes the standard deviations of intermeeting period revision contributions over the full period from 1987 to present. Over the entire sample, the standard deviation of intermeeting changes in ten year yields is about 35 basis points. The standard deviations for the two primary components—the

Table 2: Standard Deviations of Intermeeting Period Revisions: 1987-2004

Variable	**Standard Deviation (basis points)**
Ten Year Zero Coupon Yield Revision	35
Term Premium Revision Contribution	27
Expected Future Short Rate Revision Contribution	29
Unemployment Rate Revision Contribution	7
Inflation Revision Contribution	17
GDP Growth Revision Contribution	7
Policy Stance Revision Contribution	22

term premium revision contribution and the expected future short rate contribution—are both nearly 30 basis points. Among the components of the expected future short rate revision contribution, the inflation and policy stance revision contributions both have standard deviations of about 20 basis points. The standard deviation of the growth and unemployment rate revision contributions are only about 7 basis points. Thus, the interest rate accounting exercise suggests that, on average, volatility in term premiums and expected future short rates are about equally important factors underlying the volatility in long-term yields. The volatility in the expected future short rate contributions, in turn, is driven primarily by volatility in the expected future inflation and policy stance contributions.

4. Concluding Comments

While the analytical framework developed in this paper lacks the formal structure common in the no-arbitrage literature, the empirical model does have some important compensating virtues for policy economists. First and foremost, the model is rich enough to capture and quantify many of the factors that most occupy the attention of policy analysts—the impact of changes in inflation expectations, the market's assessment of the outlook for spending and economic growth, and changing perceptions of the stance of monetary policy. And the distinguishing feature of the empirical model—time variation in the perceived underlying process driving the short-rate and key macro variables—does seem to be important in understanding the evolution of the yield curve. In particular, changing perceptions of the stance of monetary policy and also the expected path for inflation emerge as key influences on the expected path for short-rates. There are, however, many yield curve puzzles that remain. Perhaps most notably, the model attributes a surprisingly large portion of the high frequency movements in longer-term yields to changes in term premiums—and this despite the fact that the very flexible framework employed in the analysis allows fluctuations in the expected paths for macro variables to play a much more prominent role in explaining yield curve movements than in more traditional models. Moreover, although the model lacks a structural specification for term premiums, explaining the volatile behavior of the estimated term premiums by relying only on the fairly smooth, downward-trending estimates of uncertainty about macro variables does not appear promising. In short, there still appears to be a very considerable measure of "excess volatility" in long-term interest rates from a macroeconomist's perspective and the issues raised twenty-five years ago by Shiller (1979) seem, if anything, all the more relevant today.

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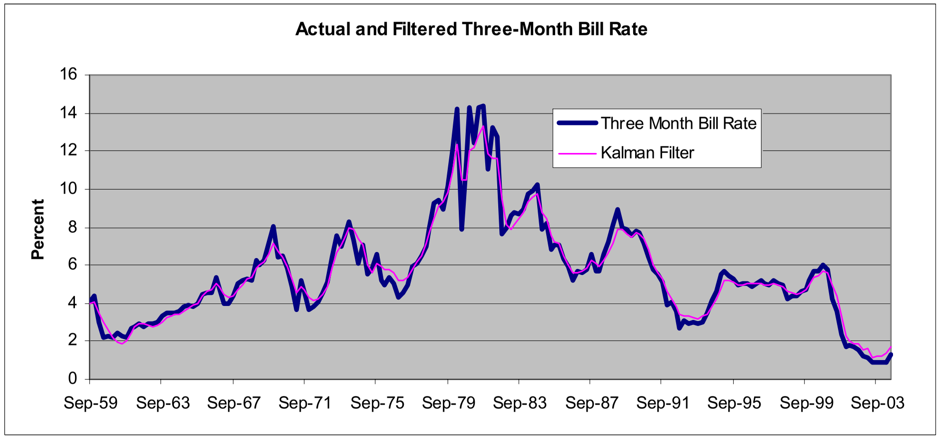
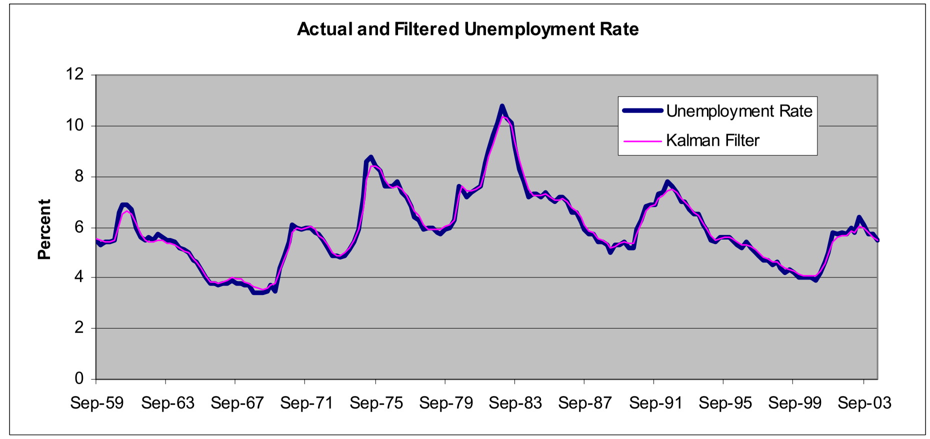
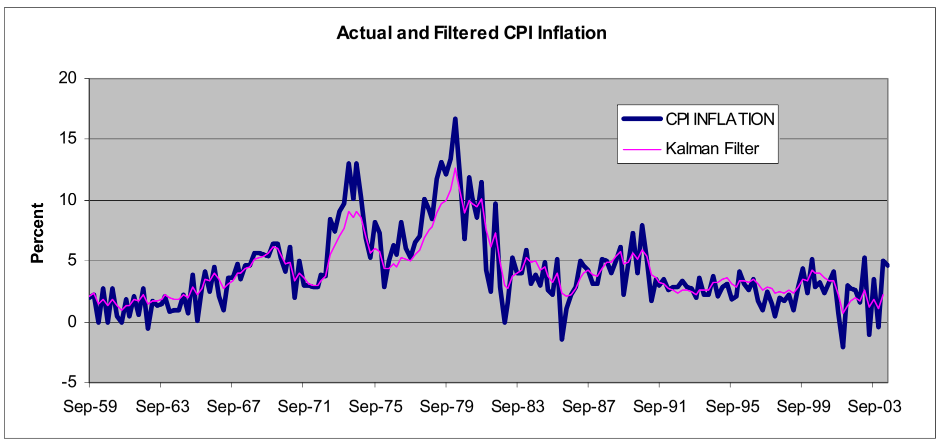
Exhibit 1: Smoothed Macro Data Series







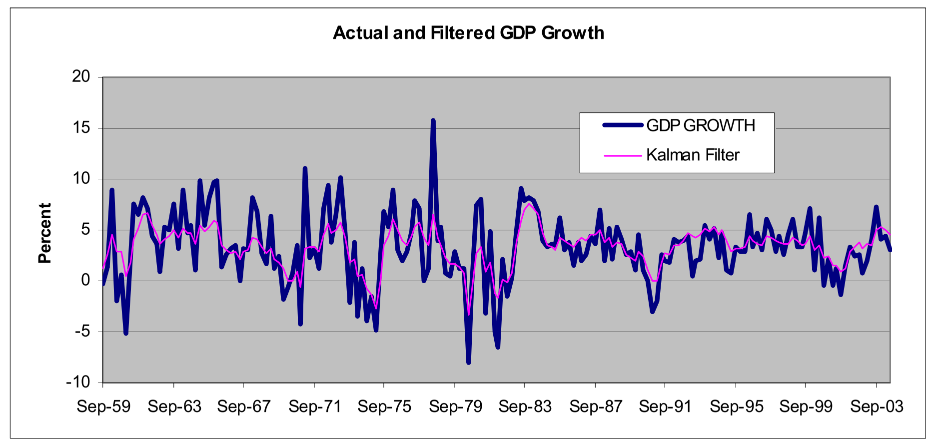


Exhibit 2: Forward Rate Decomposition

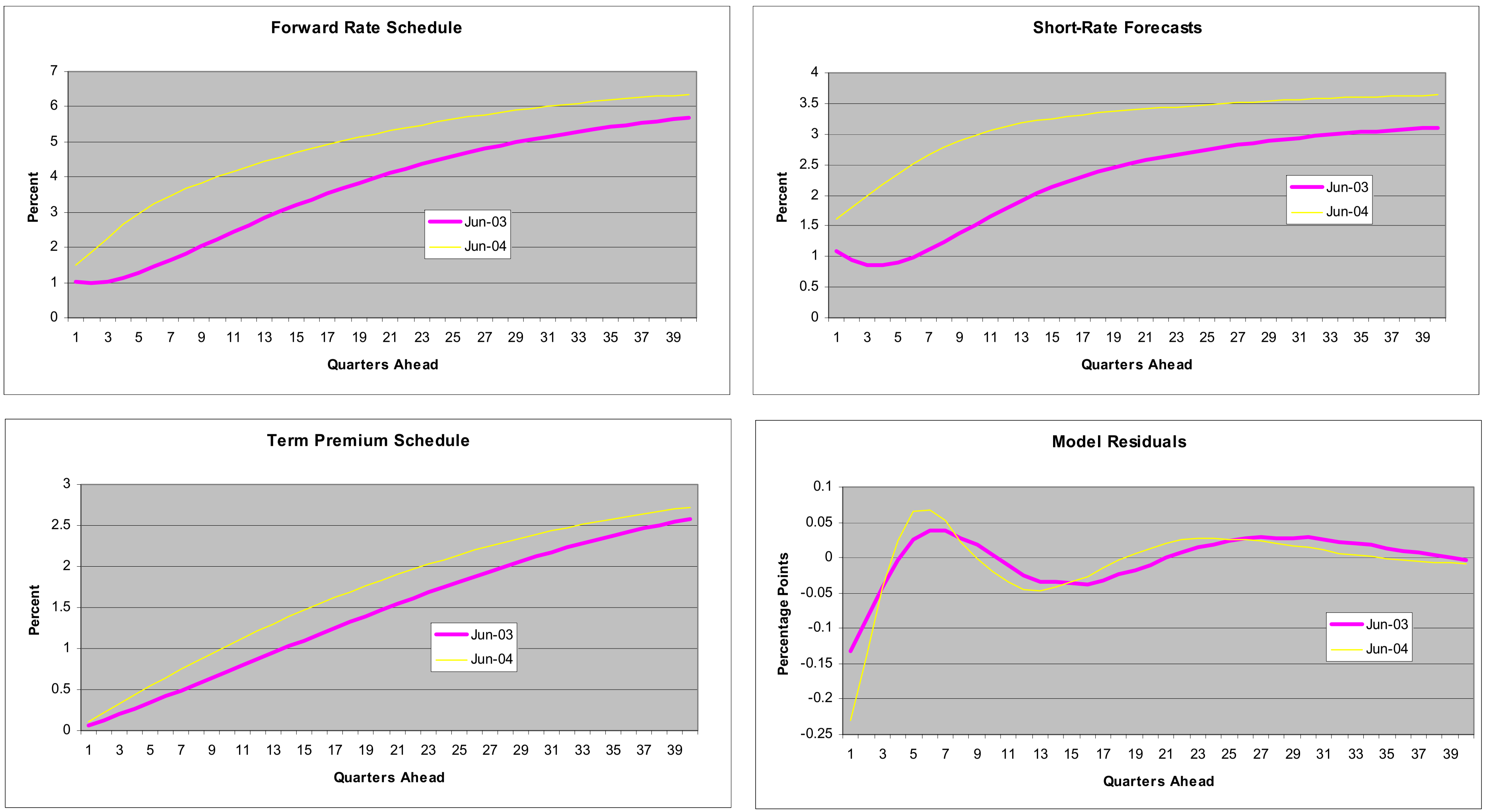


Exhibit 3: Time-Series of Forward Rate Decomposition

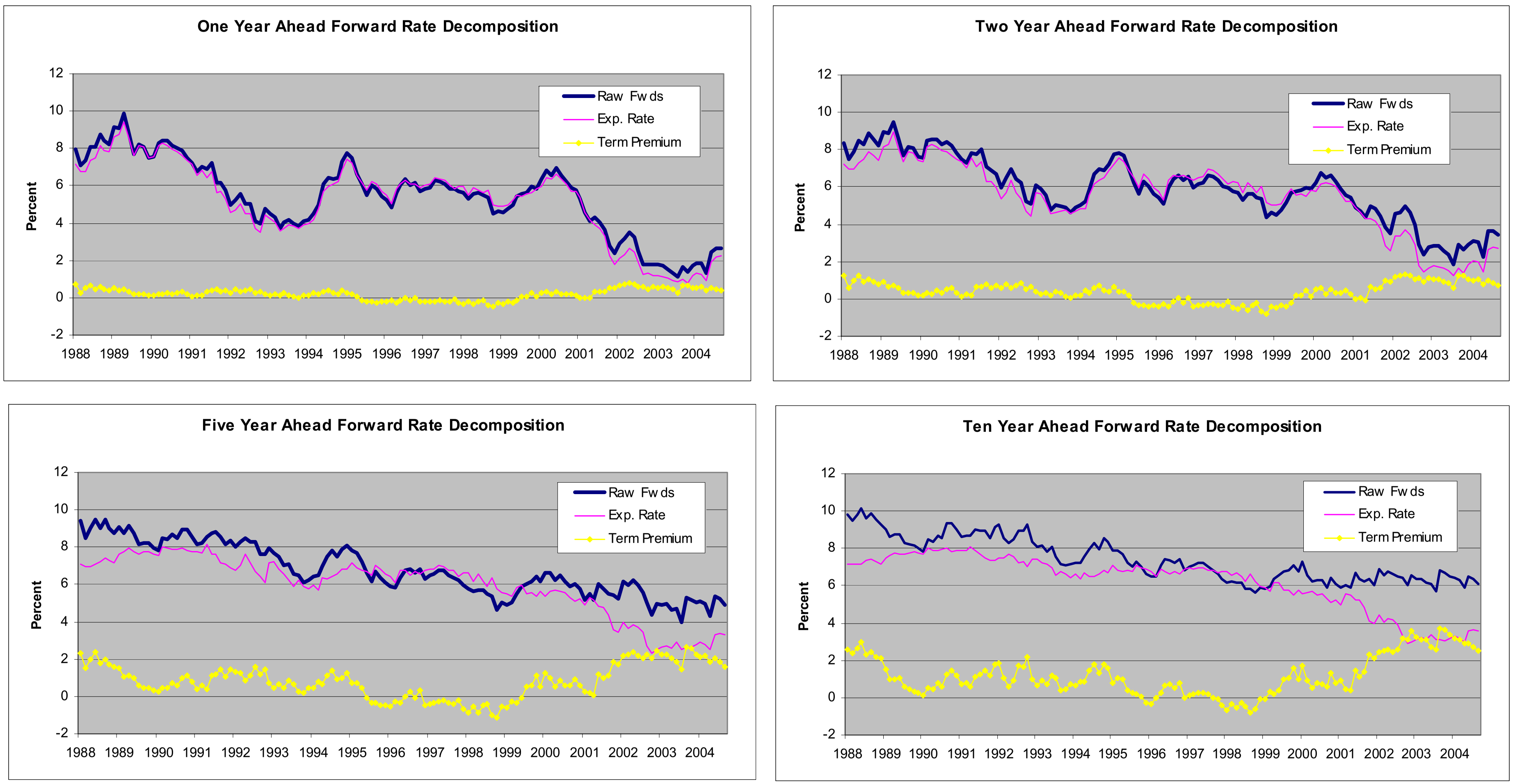
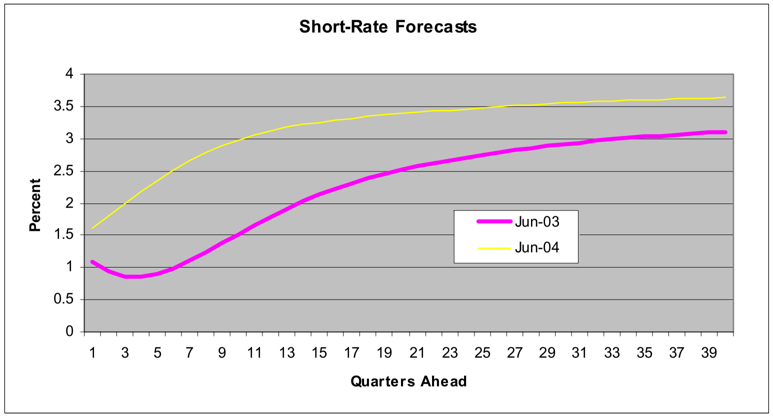
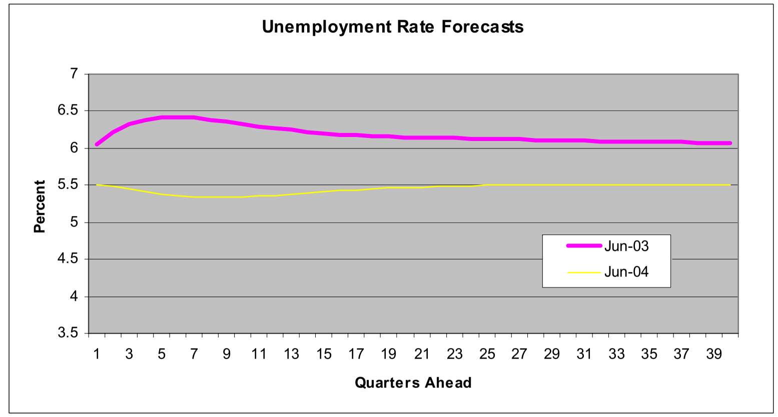
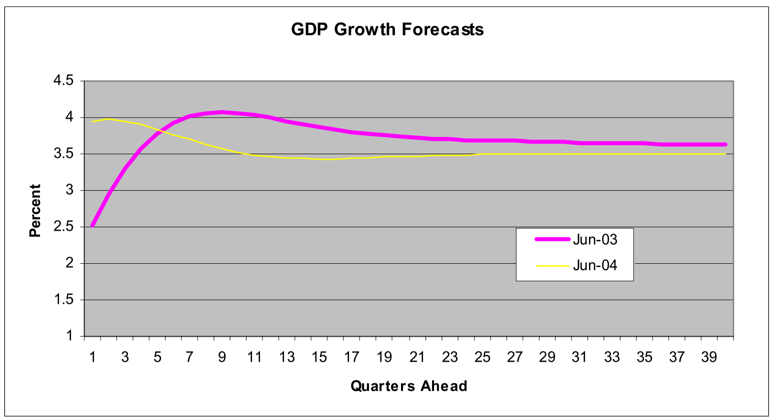
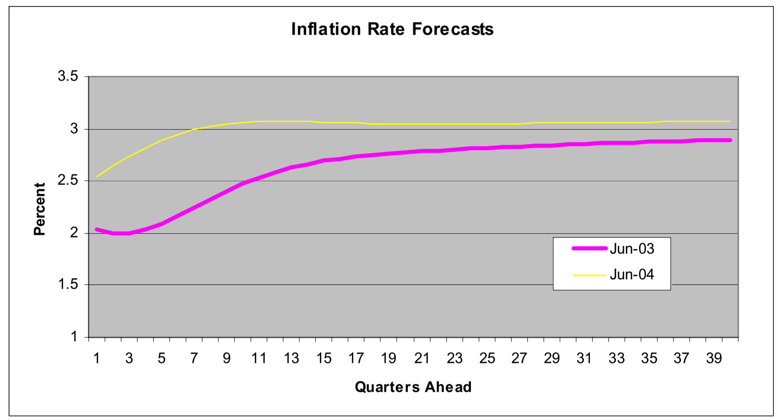


Exhibit 4: Market-Based Forecasts for Other Macro Variables









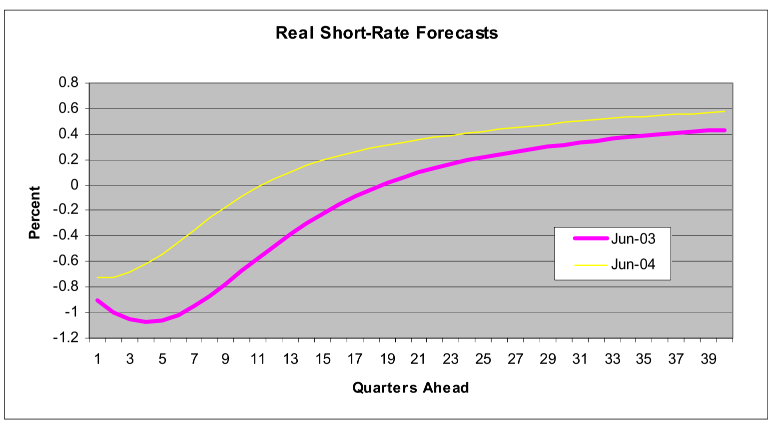
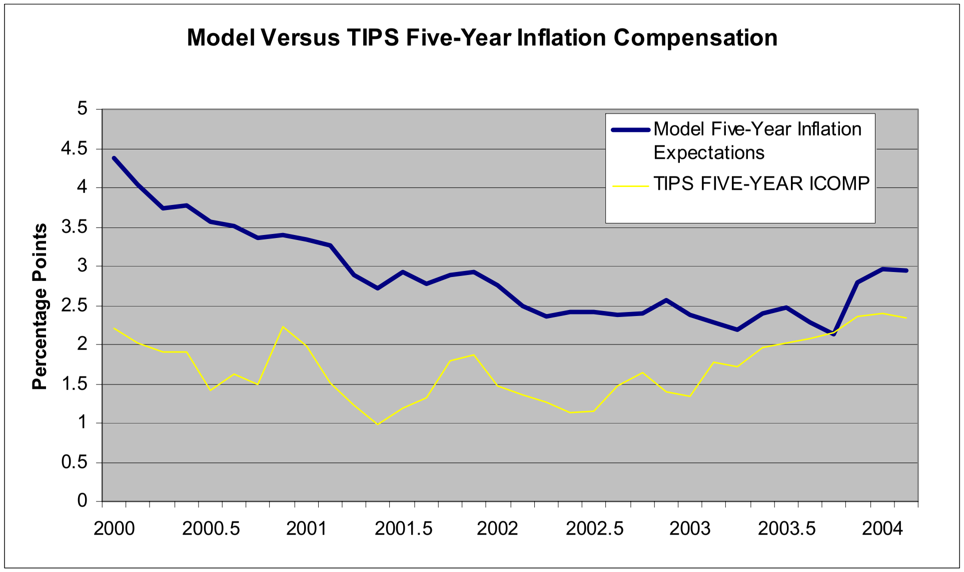


Exhibit 5: Model-Based Inflation Expectations and TIPS Inflation Compensation



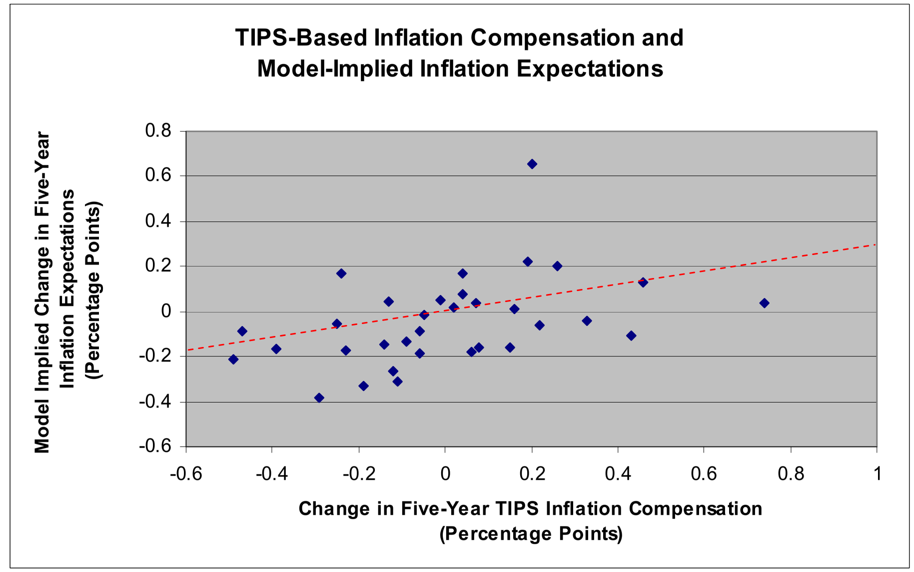


Exhibit 6: Time Series for Expected Short-Rate Decomposition

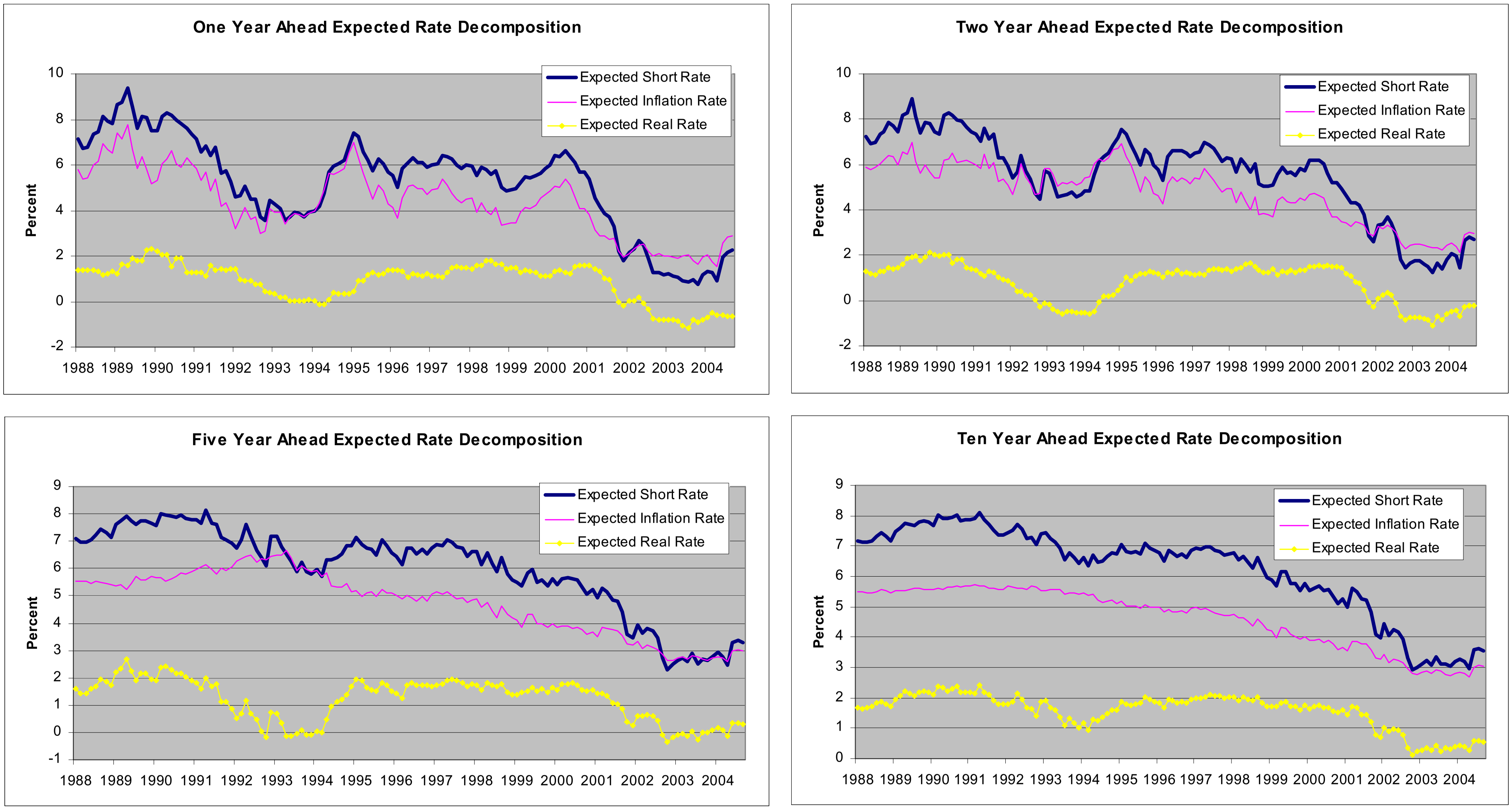
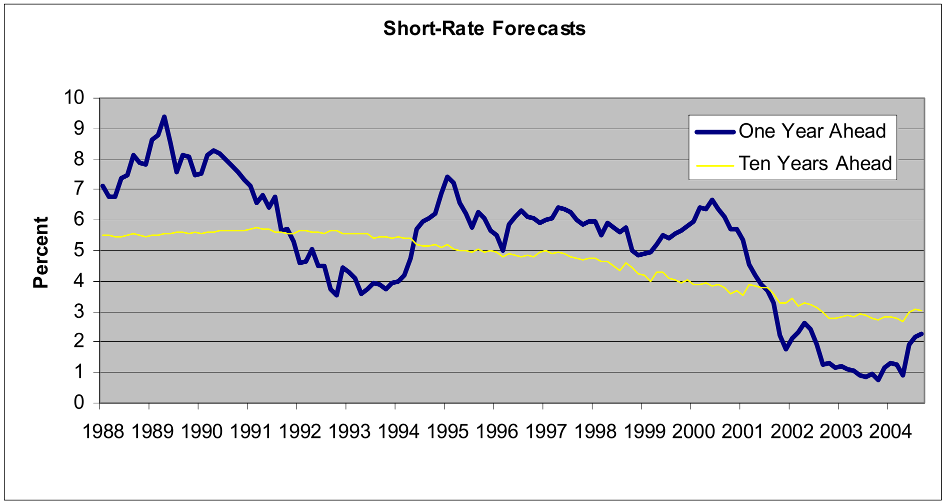
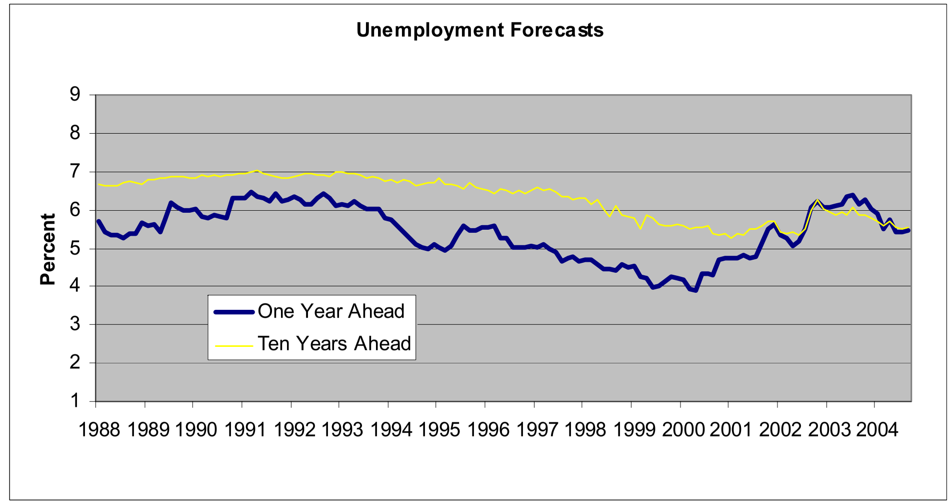
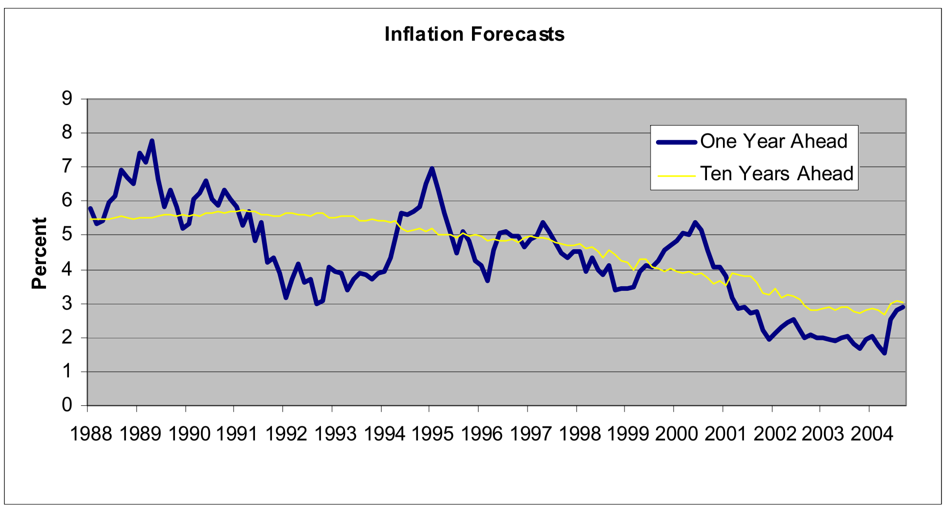


Exhibit 7: Time Series for Forecasts of Economic Variables







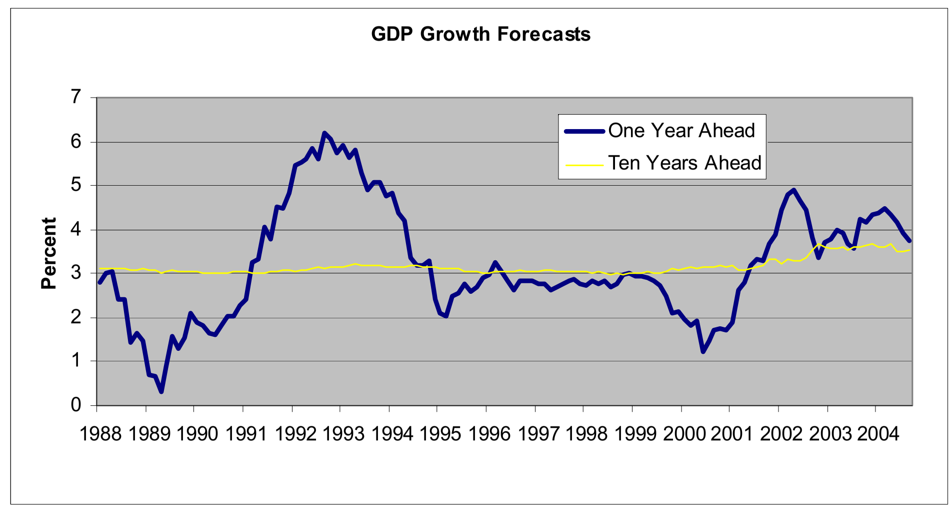
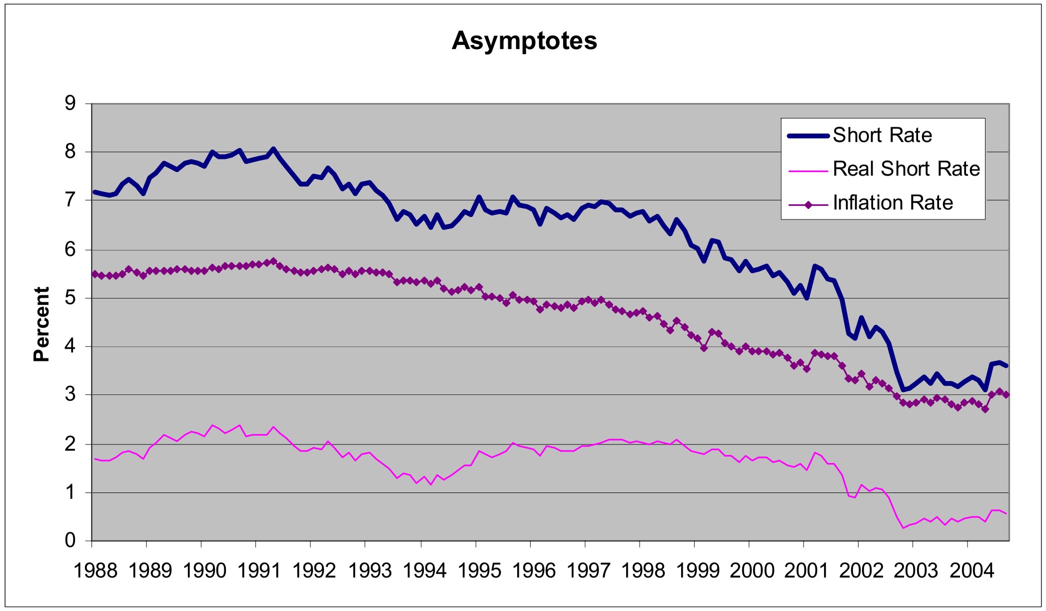


Exhibit 8: Variation in Asymptotes



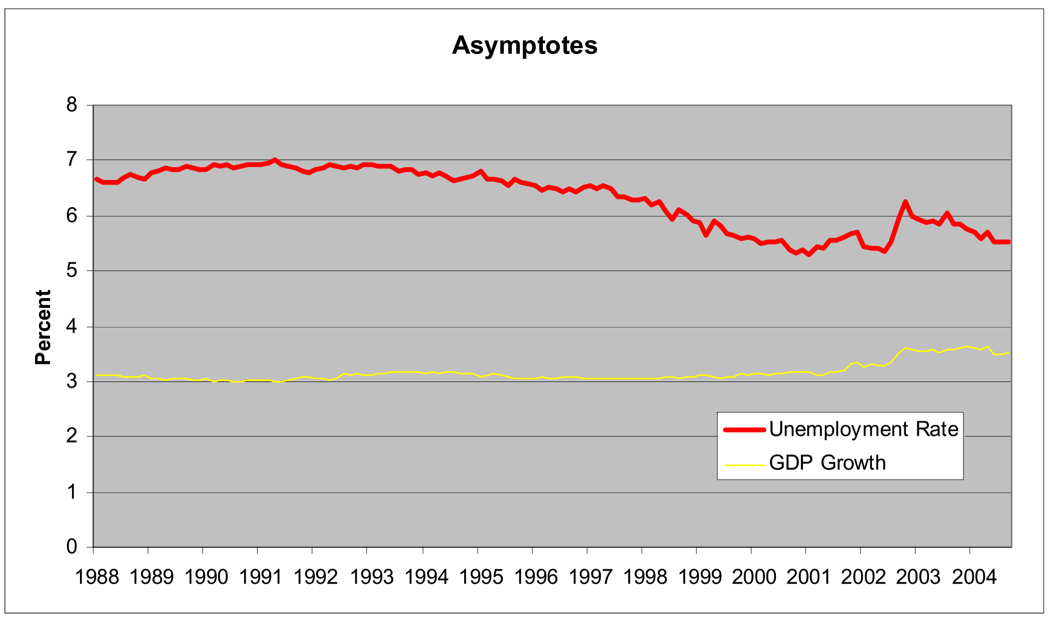
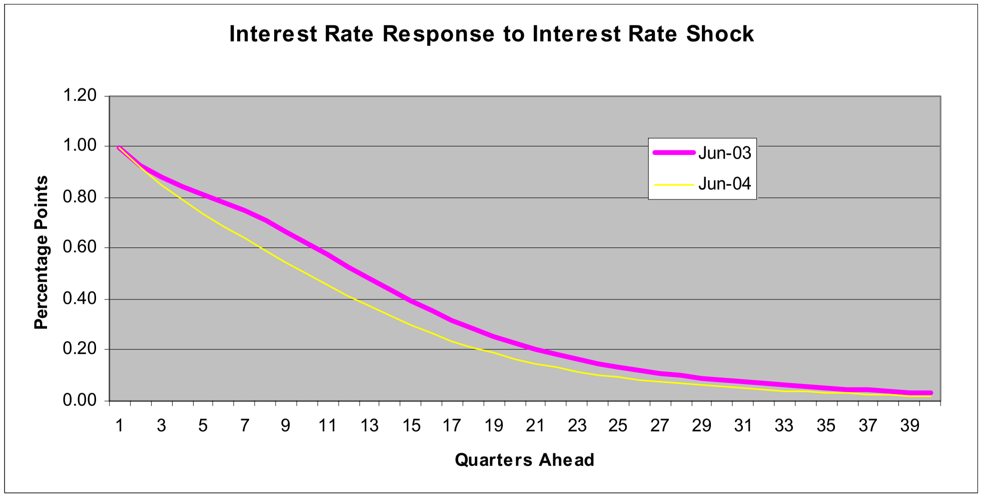
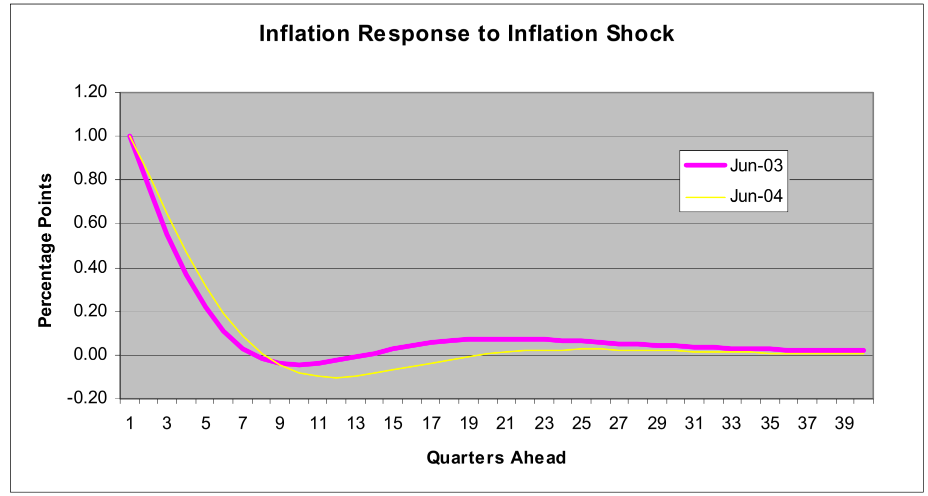
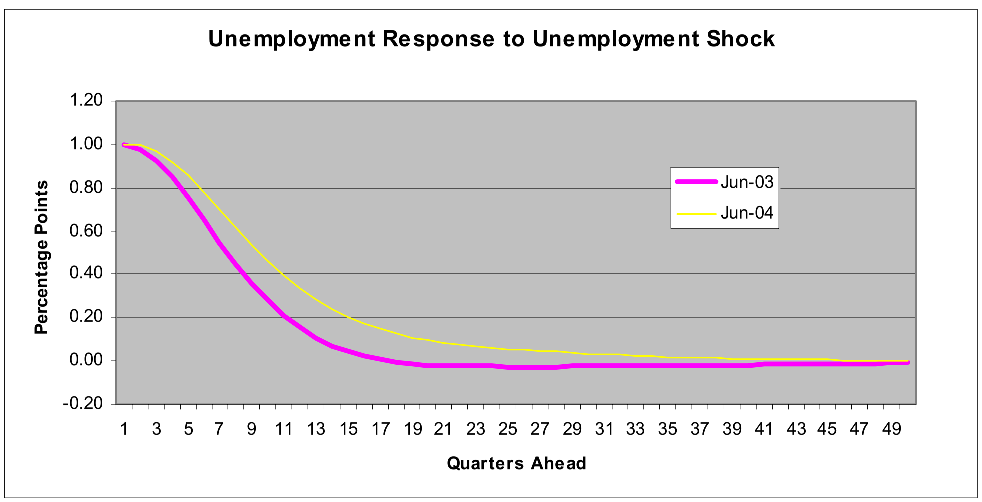


Exhibit 9: Variation in Model Dynamics







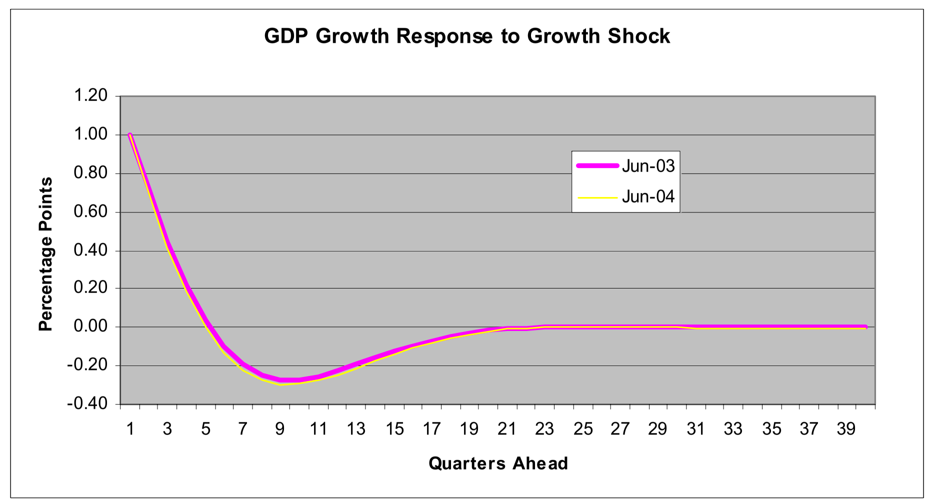


Exhibit 10: Variation in Model Dynamics

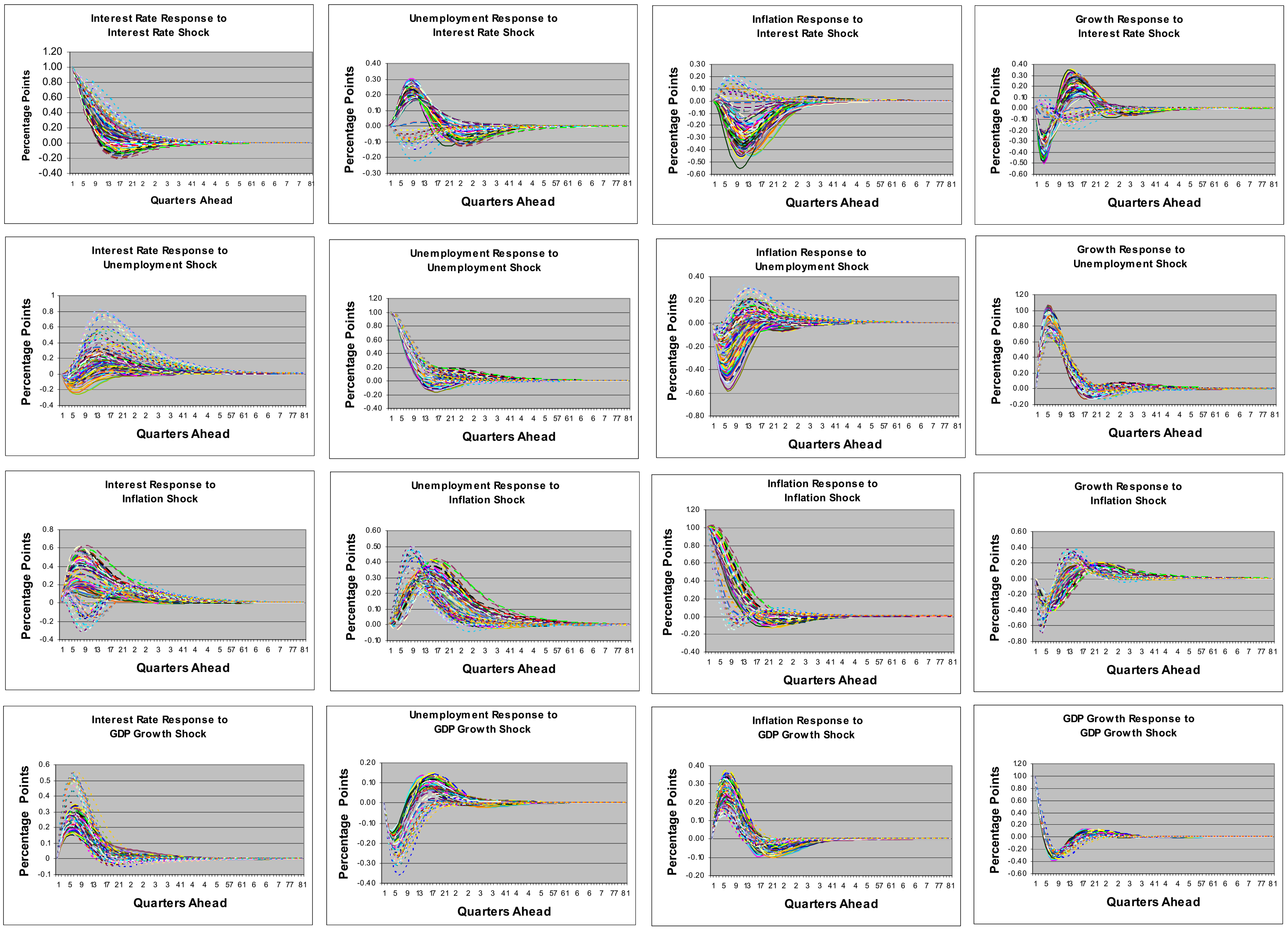
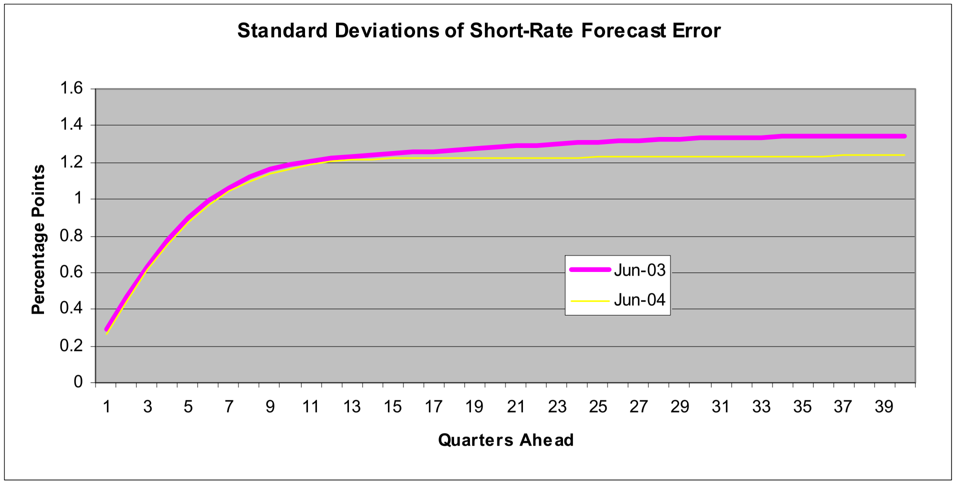
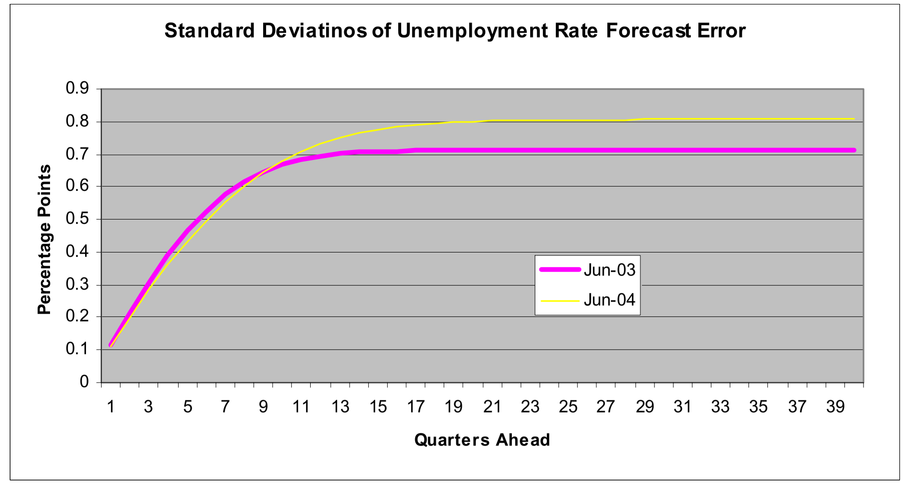
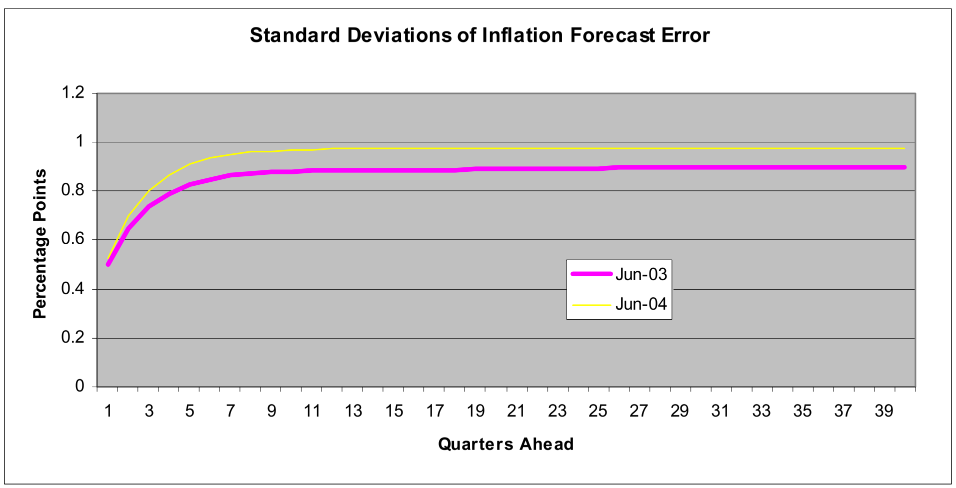


Exhibit 11: Term Structure of Volatility







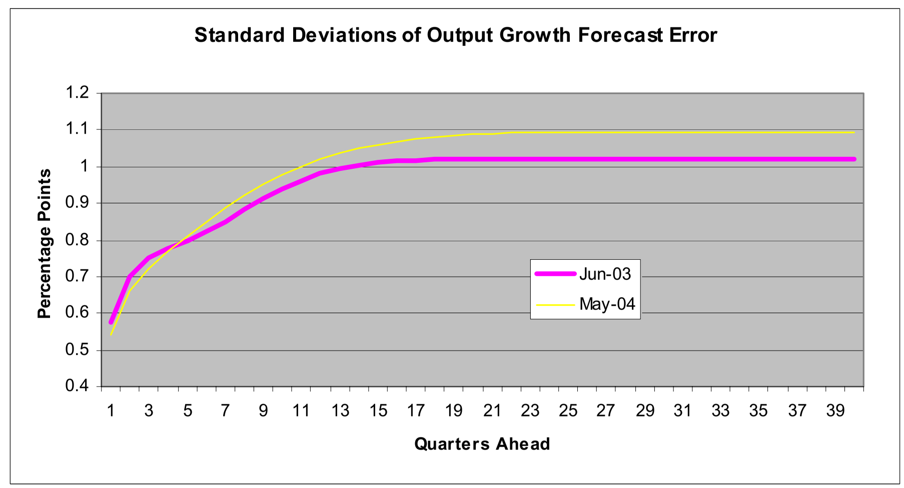
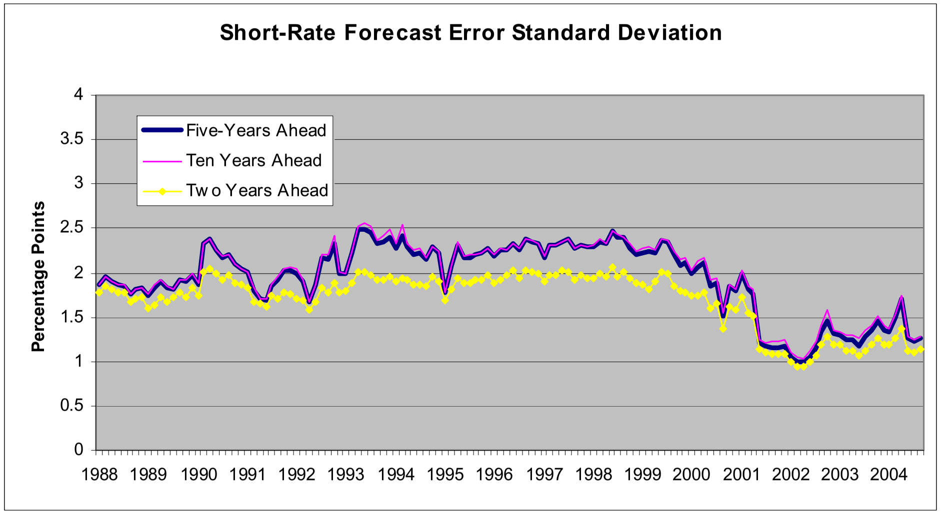
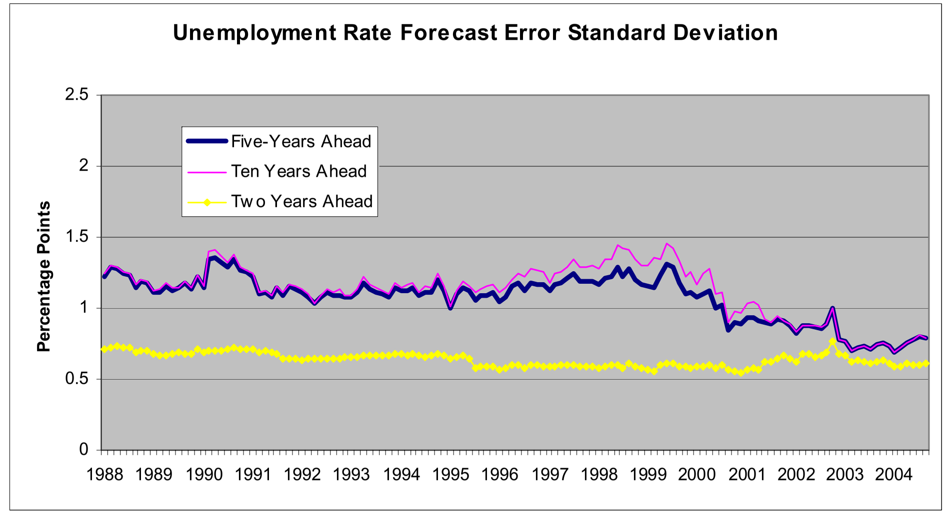
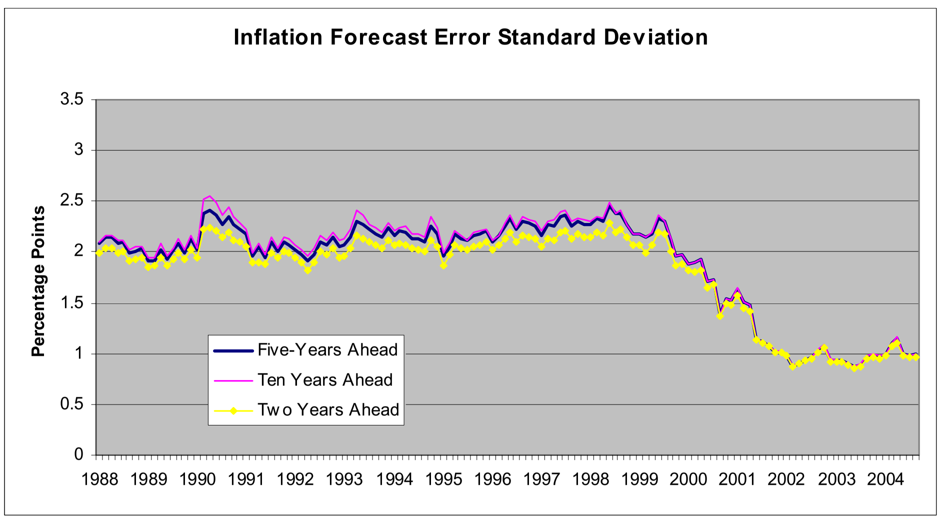


Exhibit 12: Time Series of Volatilities







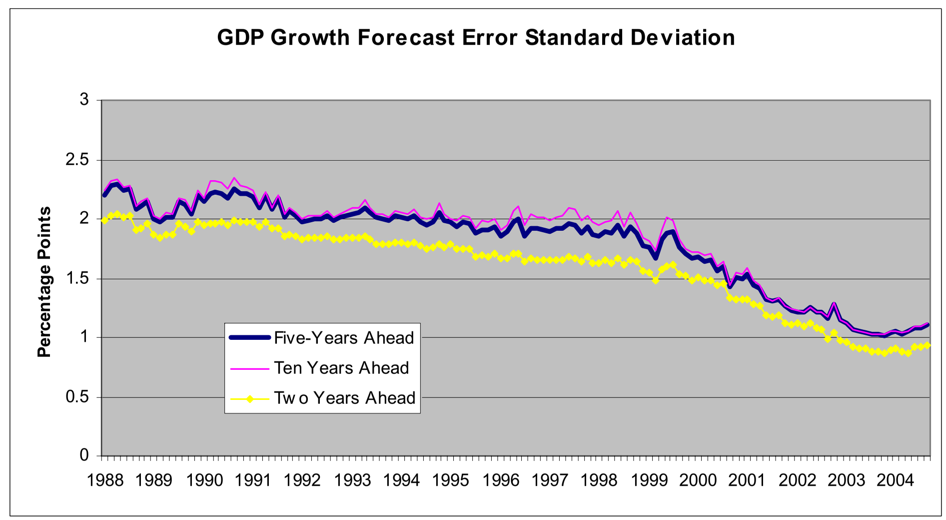


Exhibit 13: Model Shocks

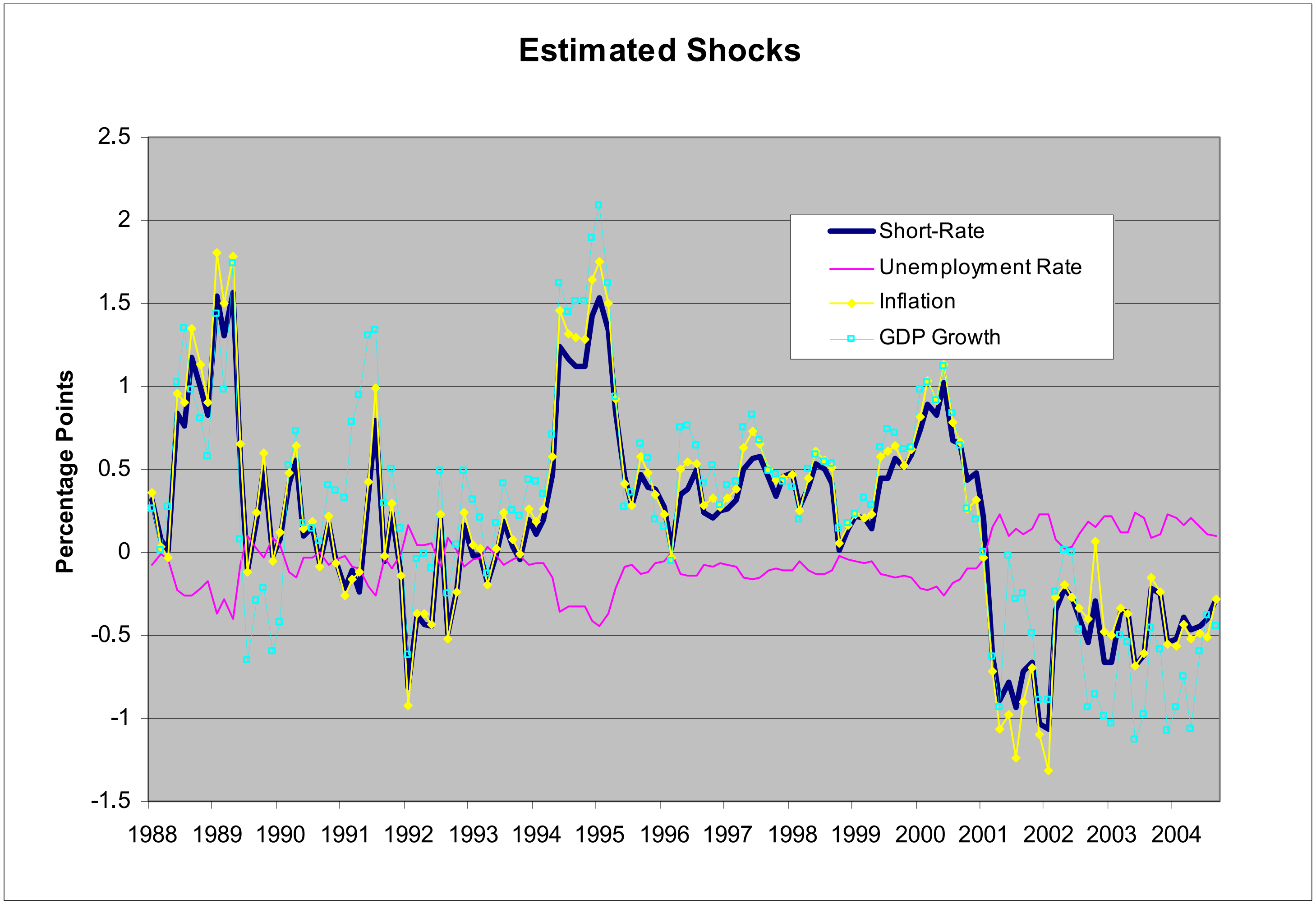
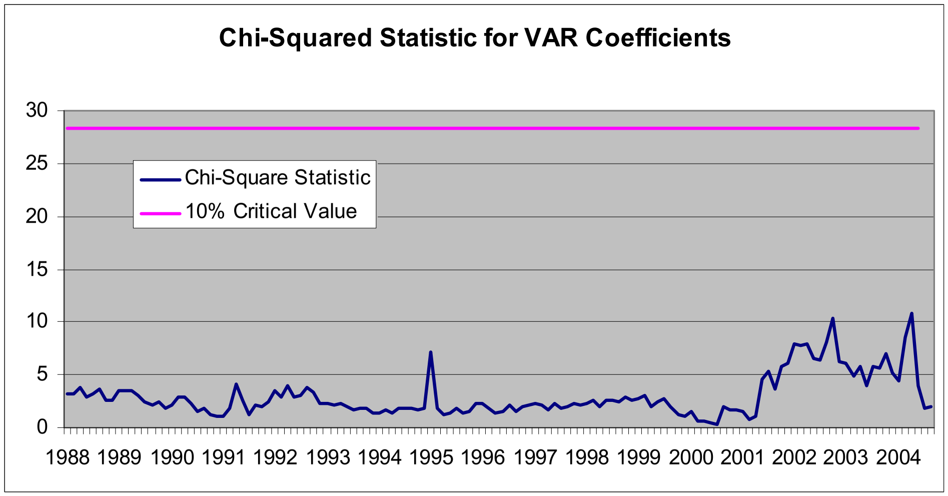


Exhibit 14: Model Diagnostics



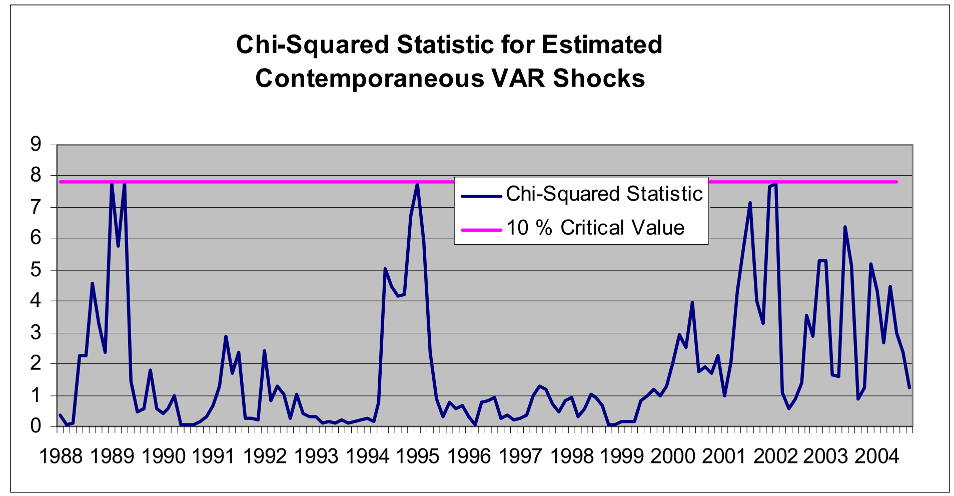
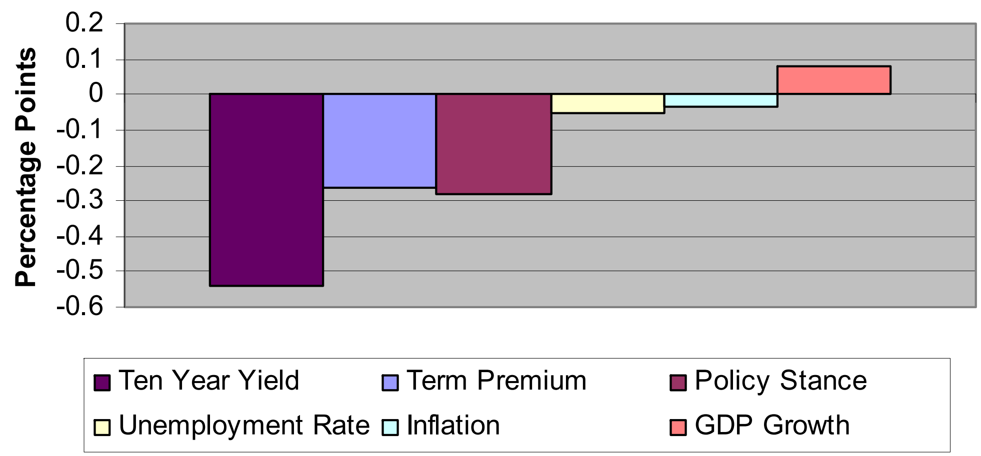


Exhibit 15

**Interest Rate Accounting:
May/June 2003 Intermeeting Period**



**Interest Rate Accounting:
June/August 2003 Intermeeting Period**