Inflation Targeting and the Anchoring of Inflation Expectations in the Western Hemisphere

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Many central banks have adopted a formal inflation-targeting framework based on the belief and the theoretical predictions that an explicit and clearly communicated numerical objective for the level of inflation over a specified period would, in itself, be a strong communication device that would help anchor long-term inflation expectations.[1] Empirically verifying the success of inflation-targeting regimes in this dimension has been difficult, however, as survey data on long-term inflation expectations tend to be of limited availability and low frequency.[2]

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1. See, for example, Leiderman and Svensson (1995); Bernanke and Mishkin (1997); Svensson (1997); Bernanke and others (1999).
2. For an analysis using semiannual survey data on long-run inflation expectations in the 1990s and early 2000s for a panel of countries, see Levin and Piger (2002).

Monetary Policy under Inflation Targeting, edited by Frederic Mishkin and Klaus Schmidt-Hebbel, Santiago, Chile.

In this paper, we use daily bond yield data for Canada, Chile, and the United States to investigate whether long-term inflation expectations in these countries are anchored, essentially extending the analysis of Gürkaynak, Sack, and Swanson (2005) and Gürkaynak, Levin, and Swanson (2006) to examine comparable data for Canada and Chile. Of these three countries, Canada and Chile have been formal inflation targeters throughout much of the 1990s and 2000s, while the United States has not had an explicit numerical inflation objective. We test the success of inflation targeting in anchoring long-term inflation expectations by comparing the behavior of long-term nominal and indexed bond yields across these three countries in response to important economic developments. Forward inflation compensation—defined as the difference between forward rates on nominal and inflation-indexed bonds—provides us with a high-frequency measure of the compensation that investors require to cover the expected level of inflation, as well as the risks associated with inflation, at a given horizon. If far-ahead forward inflation compensation is relatively insensitive to incoming economic news, then one could reasonably infer that financial market participants have fairly stable views regarding the distribution of long-term inflation outcomes. This is precisely the outcome one would hope to observe in the presence of an explicit and credible inflation target.

The daily frequency of our bond yield data, together with the frequent release of important macroeconomic statistics and monetary policy announcements, provides a large event-study data set for our analysis. This holds even for samples that span only a few years—the period for which we have inflation-indexed bond data for the United States and long-term nominal bond data for Chile. Thus, in contrast to previous empirical work using quarterly or even semiannual data, we are able to bring to bear thousands of daily observations of the response of long-term bond yields to major economic news releases in Canada, Chile, and the United States.

For the United States, we find that far-ahead forward nominal interest rates and inflation compensation respond significantly and systematically to a wide variety of macroeconomic data releases and monetary policy announcements. These responses are all consistent with a model in which the private sector's view of the central bank's long-run inflation objective is not strongly anchored, as we show. In Canada, far-ahead forward nominal interest rates and inflation compensation display little or no such sensitivity to either domestic or foreign economic news. Thus, the anchoring of long-run inflation expectations in Canada appears to be stronger than in the United States. Finally, the data for Chile is more limited in terms of the sample period, the depth and breadth of fixed income markets, and the availability of domestic macroeconomic data releases. Despite these limitations, we do not find significant responses of far-ahead inflation compensation in Chile with respect to domestic or foreign macroeconomic news.[3]

The remainder of the paper proceeds as follows. Section 1 presents two reference models of the economy to act as benchmarks for comparison with our empirical results. Section 2 investigates the responses of far-ahead forward interest rates and inflation compensation in the United States to economic news and shows that these rates respond by much more than standard models would predict. Section 3 discusses possible explanations for this finding. Section 4 repeats our empirical analysis for Canada and Chile to investigate the extent to which inflation targeting may help anchor the private sector's views regarding the long-run inflation objective of the central bank. Section 5 concludes. An appendix provides a detailed description of all the data used in our analysis.

1. Long-Run Implications of Macroeconomic Models

To aid the interpretation of our econometric results, it is useful to have a reference model as a benchmark. We consider two standard macroeconomic models: a pure new Keynesian model (taken from Clarida, Galí, and Gertler, 2000) and a modification of that model that allows for a significant fraction of backward-looking or rule-of-thumb agents (taken from Rudebusch, 2001). These two models can be thought of as different parameterizations of the following equations:

$$\pi_t = \mu_\pi E_t \pi_{t+1} + (1-\mu_\pi) A_\pi(L)\pi_t + \gamma y_t + \varepsilon_t^\pi \text{ and} \tag{1}$$

$$y_t = \mu_y E_t y_{t+1} + (1-\mu_y) A_y(L) y_t - \beta(i_t - E_t \pi_{t+1}) + \varepsilon_t^y, \tag{2}$$

where π denotes the inflation rate, y the output gap, and i the short-term nominal interest rate, and ε^π and ε^y are independent

3. Ertürk and Özlale (2005) obtain a similar finding of anchored expectations for Chile using a GARCH specification on monthly Chilean data.

and identically distributed (i.i.d.) shocks.[4] The parameters μ_π and μ_y describe the degree of forward-looking behavior in the model, and the lag polynomials $A_\pi(L)$ and $A_y(L)$ summarize the parameters governing the dynamics of any backward-looking components of the model.

The two models differ in the extent of their forward-looking behavior. The pure new Keynesian model assumes that agents are completely forward looking ($\mu_\pi = \mu_y = 1$), and the parameter values for the equations are taken from Clarida, Galí, and Gertler (2000). A number of authors, however, estimate much smaller values of μ_π (around 0.3) to match the degree of inflation persistence observed in U.S. data (for example, Fuhrer, 1997; Roberts, 1997; Rudebusch, 2001; Estrella and Fuhrer, 2002). Thus, in the second model considered, we set $\mu_\pi = 0.3$ and take parameter values from Rudebusch (2001).[5] Note that Rudebusch's model is among the most persistent of the hybrid new Keynesian models in the literature, owing to the inclusion of several lags of output and inflation in equations (1) and (2) and a particularly low value of μ_y (Rudebusch assumes $\mu_y = 0$) in the income-spending (IS) equation (equation 2).

We close these two models with an interest rate rule of the following form:

$$i_t = (1-c)\left[(1+a)\bar{\pi}_t + by_t\right] + ci_{t-1} + \varepsilon_t^i, \tag{3}$$

where $\bar{\pi}$ denotes the trailing four-quarter moving average of inflation, ε^i is an i.i.d. shock, and a, b, and c are the parameters of the rule.[6] Note that the policy rule is both backward-looking, in that the interest rate responds to current values of the output gap and inflation rather than their forecasts, and inertial, in that it includes the lagged federal funds rate. Both of these characteristics tend to add inertia to the short rate, which, together with the persistence of the Rudebusch model, generally gives the model the best possible chance to explain the term structure evidence we find below. We include an interest rate shock, ε_t^i, for the purpose of generating impulse response functions.

4. These variables are all normalized to have steady-state values of zero.

5. Rudebusch estimates and uses a value of $\mu = 0.29$ in the inflation equation and sets $\mu = 0$ in the output equation, so we use those values as well. There are also some minor timing differences between the specification of Rudebusch's model and equations (1) and (2). To generate the impulse response functions in figure 1, we use the model exactly as specified in Rudebusch (2001), but these differences in specification have no discernible effect on our results.

6. We use the values of a, b, and c estimated by Rudebusch (2002) from 1987:4 to 1999:4: namely, $a = 0.53$, $b = 0.93$, and $c = 0.73$.

The three panels of figure 1 show the response of the short-term nominal interest rate to a one-percent shock to the inflation equation, the output equation, and the interest rate equation, respectively, under our two baseline models.[7] In the pure new Keynesian (Clarida, Galí, and Gertler) model, the effect of the macroeconomic and monetary policy shocks on the short-term interest rate dies out very quickly, generally within a year. The interest rate displays much more persistence in the partially backward-looking (Rudebusch) model. Even in that model, however, the short-term interest rate essentially returns to its steady-state level well within ten years after each shock.

2. The Sensitivity of U.S. Long-Term Interest Rates to Economic News

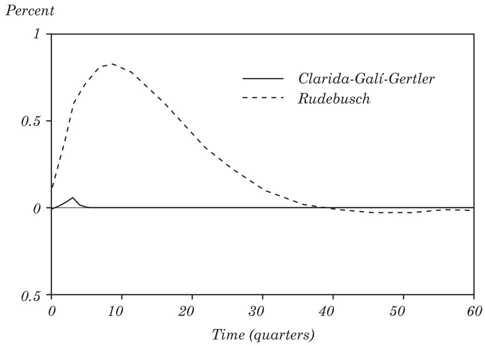
We now turn to how well the above model predictions are matched by U.S. data. The models predict that macroeconomic data releases and monetary policy announcements should affect the path of nominal interest rates only in the short run. To examine whether the U.S. data match the predictions of the models, we must look beyond the response of interest rates in the first few years after a shock and instead focus on the behavior of forward interest rates several years ahead.

Forward rates are often a very useful means of interpreting the term structure of interest rates. For a bond with a maturity of m years, the yield $r_t^{(m)}$ represents the rate of return that an investor requires to lend money today in return for a single payment m years in the future (for the case of a zero-coupon bond). By comparison, the k-year-ahead one-year forward rate $f_t^{(k)}$ represents the rate of return from period $t + k$ to period $t + k + 1$ that the same investor would require to commit today to a one-year loan beginning at time $t + k$ and maturing at time $t + k + 1$. The linkage between these concepts

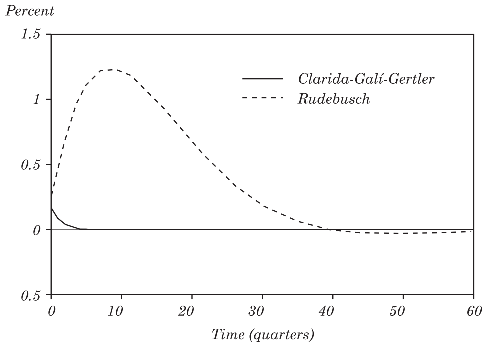
7. In a discussion of our paper at the Central Bank of Chile, Eric Parrado reported impulse response functions using the small open economy international macroeconomic model of Galí and Monacelli (2005), roughly calibrated to match the data in Canada and Chile. The results from those impulse response functions were consistent with our analysis for the standard closed economy new Keynesian models presented here: in particular, short-term interest rates returned to steady state well within ten years of a shock. Indeed, that model returned to steady state even more quickly—within just four or five years, compared to seven or eight years for the Rudebusch model. We believe this difference is due to the persistent parameters of the Rudebusch model, rather than to the lack of an open economy transmission mechanism in that model.

Figure 1. Impulse Response Functions for Standard Macroeconomic Models

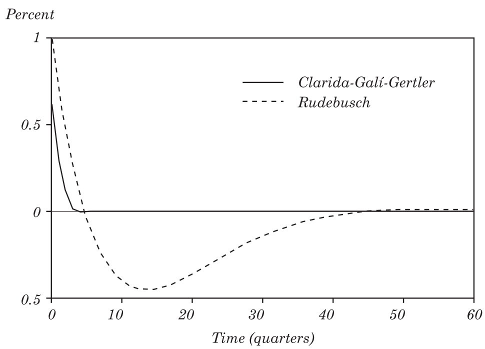
Interest rate response to a 1 percent inflation shock



Interest rate response to a 1 percent output shock



Interest rate response to a 1 percent interest rate shock



Source: Authors' computations.

is simple: an m-year (continuously-compounded) zero-coupon security can be viewed as a sequence of one-year forward agreements over the next m years:[8]

$$f_t^{(k)} = (k+1)r_t^{(k+1)} - kr_t^{(k)}. \tag{4}$$

For our analysis, we use Federal Reserve Board data on forward interest rates for U.S. Treasury securities.[9] Given our interest in measuring long-term expectations, our analysis focuses on the longest maturity for which we have high-quality bond yield data. The liquidity and breadth of the markets for government securities at and around the ten-year horizon thus lead us to focus on the one-year forward rate nine years ahead (that is, the one-year forward rate ending in ten years). The analysis of the previous section shows that this horizon is sufficiently far out for standard macroeconomic models to largely return to their steady states, so that any movements in forward interest rates or inflation compensation at these horizons should not be due to transitory responses of the economy to an economic shock.

To measure the effects of macroeconomic data releases on interest rates, the unexpected (or surprise) component of each macroeconomic data release must be computed, since the expected component of macroeconomic data releases should have no effect in forward-looking financial markets.[10] Using the surprise components of macroeconomic data releases, where expectations are measured just a few days before the actual release, also removes any possible issue of endogeneity arising from interest rates feeding back to the macroeconomy. Any such effects, to the extent that they are systematic or predictable, will be incorporated into the market forecast for the statistical release.

To measure the surprise component of each data release, we compute the difference between the actual release and the median forecast of

8. If we could observe zero-coupon yields directly, computing forward rates would be as simple as this. In practice, however, most government bonds in the United States and abroad make regular coupon payments, and thus the size and timing of the coupons must be accounted for to translate observed yields into the implied zero-coupon yield curve. In the results presented below, we also investigate whether the use of U.S Treasury STRIPS (which are zero-coupon securities that thus do not require fitting a yield curve first) alters the estimated response of far-ahead forward nominal rates in the United States. We find that the STRIPS data yield essentially identical results.

9. The Federal Reserve Board computes implied zero-coupon yields from observed, off-the-run U.S. Treasury yields using the extension of Nelson-Siegel described in Svensson (1994). Details are available in Gürkaynak, Sack, and Wright (2005).

10. Kuttner (2001) tests and confirms this hypothesis for the case of monetary policy announcements.

that release made by professional forecasters just a few days prior to the release date. For the United States, we use data on professional forecasts of the next week's statistical releases, published every Friday by Money Market Services for thirty-nine different macroeconomic data series.[11] Not all thirty-nine of these macroeconomic statistics have a significant impact on interest rates, even at the short end of the yield curve. Thus, to conserve space and reduce the number of exogenous variables in our regressions, we restrict our attention to the macroeconomic variables that Gürkaynak, Sack, and Swanson (2005) identify as having statistically significant effects on the one-year Treasury bill rate over the 1990–2002 period: capacity utilization, consumer confidence, the core consumer price index (CPI), the employment cost index (ECI), the advance (that is, first) release of real GDP, initial claims for unemployment insurance, the National Association of Purchasing Managers (NAPM) / Institute for Supply Management (ISM) survey of manufacturing activity, new home sales, employees on nonfarm payrolls, retail sales, and the unemployment rate.[12]

As with macroeconomic data releases, we must compute the surprise component of monetary policy announcements in each of our countries in order to measure the effects of these announcements on interest rates. We measure monetary policy surprises for the United States using federal funds futures rates, which provide high-quality, virtually continuous measures of market expectations for the federal funds rate (Krueger and Kuttner, 1996; Rudebusch, 1998; Brunner, 2000).[13] The federal funds futures contract for a given month settles at the end of the month based on the average federal funds rate that was realized over the course of that month. Thus, daily changes in the current-month futures rate reflect revisions to the market's expectations for the federal funds rate over the remainder of the month. As explained in Kuttner (2001) and Gürkaynak, Sack, and Swanson (2002), the

11. Several authors find the Money Market Services data to be of high quality (for example, Balduzzi, Elton, and Green, 2001; Andersen and others, 2003; Gurkaynak, Sack, and Swanson, 2005).

12. In addition to these eleven variables, Gürkaynak, Sack, and Swanson (2005) also included leading indicators and the core producer price index in their analysis. We originally included these two variables as well, but they never entered significantly into any of our regressions at even the shortest horizon at even the ten percent level. We therefore omit them from the results below to save space and reduce the number of explanatory variables. Nonetheless, our results are essentially identical whether we include these additional variables in the regressions or not.

13. Gürkaynak, Sack, and Swanson (2002) show that, among the many possible financial market instruments that potentially reflect expectations of monetary policy, federal funds futures are the best predictor of future policy actions.

change in the current month's contract rate on the day of a Federal Open Market Committee (FOMC) announcement, scaled up to account for the timing of the announcement within the month, provides a measure of the surprise component of the FOMC decision.[14] We compute the surprise component associated with every FOMC meeting and intermeeting policy action by the FOMC over our sample.[15]

2.1 The Sensitivity of U.S. Interest Rates to Economic News

Table 1 reports results for nominal interest rates in the United States over the 1994–2005 period.[16] Each column provides results from a regression of daily changes in the corresponding interest rate on the surprise component of the macroeconomic data releases and monetary policy announcements listed at the left.[17] We regress the change in interest rates on all of our macroeconomic and monetary policy surprises jointly to properly account for days on which more than one piece of economic news was released. To facilitate interpreting our coefficient estimates, we normalize each macroeconomic surprise by its standard deviation. Each coefficient in the table thus estimates the interest rate response in basis points per standard deviation surprise in the corresponding macroeconomic statistic. The one exception to this rule is the monetary policy surprises, which we leave in basis points, so that these coefficients represent a basis point per basis point response.

14. To avoid very large scale factors, if the monetary policy announcement occurs in the last seven days of the month, we use the next-month contract rate instead of scaling up the current-month contract rate.

15. The only exception is that we exclude the intermeeting 50 basis point easing on 17 September 2001, because financial markets were closed for several days prior to that action and because that easing was a response to a large exogenous shock to the U.S. economy and financial markets. We would thus have difficulty disentangling the effect of the monetary policy action from the effect of the shock itself on financial markets that day.

16. Our STRIPS data begin in 1994, so we restrict analysis in table 1 to the post-1994 period. Gürkaynak, Sack, and Swanson (2005) report very similar results for the 1990–2002 period using forward rates from a fitted yield curve.

17. Although we have almost one thousand daily observations in each of these regressions, most of the elements of any individual regressor are zero, because any given macroeconomic statistic is only released once a month (or once a quarter in the case of GDP and once a week in the case of initial claims). We restrict attention in all our regressions to those days on which some macroeconomic statistic was released or a monetary policy announcement was made, but our results are not sensitive to this restriction.

Table 1. U.S. Forward Rate Responses to Domestic Economic News, 1994–2005[a]

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward nominal rate ending in ten years, from STRIPS*
Capacity utilization	1.76***	1.24**	0.80
	(3.78)	(2.05)	(1.21)
Consumer confidence	1.36***	1.04*	0.88
	(3.13)	(1.85)	(1.43)
Core CPI	1.92***	1.47*	1.80**
	(3.29)	(1.94)	(2.16)
Employment cost index	1.66**	1.87*	1.24
	(2.28)	(1.98)	(1.20)
Real GDP (advance)	1.37*	0.36	–0.08
	(1.95)	(0.40)	(–0.08)
Initial jobless claims	–0.91***	–0.59**	–0.62**
	(–4.16)	(–2.07)	(–2.00)
NAPM/ISM manufacturing survey	2.40***	2.54***	2.79***
	(5.58)	(4.55)	(4.56)
New home sales	0.77*	0.85	1.01*
	(1.88)	(1.60)	(1.73)
Nonfarm payrolls	4.63***	2.51***	2.62***
	(10.24)	(4.28)	(4.08)
Retail sales (excl. cars)	2.15***	1.69**	1.36*
	(3.75)	(2.26)	(1.66)
Unemployment rate	–1.63***	0.38	–0.52
	(–3.32)	(0.60)	(–0.74)
Monetary policy	0.30***	–0.17**	–0.24***
	(4.78)	(–2.14)	(–2.71)
No. observations	1,371	1,371	1,371
R^2	0.16	0.06	0.05
Joint test *p* value	0.000***	0.000***	0.000***

Source: Authors' computations.
* Statistically significant at the 10 percent level. ** Statistically significant at the 5 percent level. *** Statistically significant at the 1 percent level.
a. The sample is from January 1994 to October 2005, at daily frequency on the dates of macroeconomic and monetary policy announcements. Regressions also include a constant, a Y2K dummy that takes on the value of 1 on the first business day of 2000, and a year-end dummy that takes on the value of 1 on the first business day of any year (coefficients not reported). Macroeconomic data release surprises are normalized by their standard deviations, so these coefficients represent a basis point per standard deviation response. Monetary policy surprises are in basis points, so these coefficients represent a basis point per basis point response. Joint test *p* value is for the hypothesis that all coefficients (other than the constant and dummy variables) are zero. *t*-statistics are reported in parentheses.

The first column of table 1 reports the responses of the one-year Treasury spot rate to the economic releases as a benchmark for comparison. As one might expect from a Taylor-type rule or from casual observation of U.S. financial markets, interest rates at the short end of the term structure exhibit highly significant responses to surprises in macroeconomic data releases and monetary policy announcements. Moreover, these responses are generally consistent with what one would expect from a Taylor-type rule: upward surprises in inflation, output, or employment lead to increases in short-term interest rates, and upward surprises in initial jobless claims (a countercyclical economic indicator) cause short-term interest rates to fall. The magnitudes of these estimates seem reasonable, with a two-standard-deviation surprise leading to about a 3 to 10 basis point change in the one-year rate (depending on the statistic) on average over our sample. Monetary policy surprises lead to about a one-for-three or one-for-two response of the one-year yield to the federal funds rate. This is consistent with the view that a surprise change in the federal funds rate is often not a complete surprise to markets, but rather a moving forward or pushing back of policy changes that were already expected to have some chance of occurring in the future.

The middle column of table 1 shows the response of far-ahead forward interest rates in the U.S. to economic news. If ten years is a sufficient amount of time for the U.S. economy to return largely to steady state following an economic shock, as our simulations above suggest, and if long-term inflation expectations were firmly anchored in the United States, then one would expect to see little or no response of these rates to economic news. This is not the case, however: far-ahead forward nominal rates in the United States respond significantly to nine of the twelve macroeconomic data releases we consider, often with a very high degree of statistical significance, and a test of the joint hypothesis that all coefficients in the regression are zero is rejected with a p value on the order of 10^{-10}. Not only are the estimated coefficients statistically significant, but their magnitudes are large, often more than half as large as the effect on the short-term interest rate. Finally, the signs of these coefficients are not random, but rather they closely resemble the effect on short-term interest rates and the short-term inflation outlook. This resemblance is consistent with markets expecting some degree of pass-through of short-term inflation to the long-term inflation outlook. The case of monetary policy surprises offers perhaps the most striking example of this pattern: the estimated effect of monetary policy surprises on far-ahead nominal interest rates is opposite to the effect of surprises on the one-year spot rate—that is, a surprise monetary policy tightening causes far-ahead forward nominal rates to fall. This result echoes the finding by Gürkaynak, Sack, and Swanson (2005) for their 1990–2002 and 1994–2002 samples. It is also consistent with financial markets expecting a pass-through of the short-term inflation outlook to long-term inflation, as we demonstrate in section 3, below.

The right-hand column of table 1 reports a robustness check on the above results, in which we computed the response of the one-year forward rate ending in ten years using U.S. Treasury STRIPS rather than the Federal Reserve's smoothed yield curve data.[18] STRIPS are pure zero-coupon securities whose yields provide a direct, market-based measure of forward rates that does not require any yield curve fitting or smoothing. (On the other hand, STRIPS are less liquid than Treasury notes and bonds and thus suffer from larger bid-ask spreads and trading costs, making observed prices a less clean measure of the true shadow value of the securities and introducing some noise into our estimates.) The results in the right-hand column of table 1 are very much in line with those from the middle column: seven of the twelve macroeconomic news releases we consider lead to significant responses of ten-year-ahead forward interest rates, with estimated magnitudes that are very similar to those from our yield-curve-based estimates, and the joint hypothesis that all coefficients are equal to zero is likewise rejected at extremely high levels of statistical significance (p value on the order of 10^{-9}). All of these observations suggest that our results are not due to any artifact of yield-curve fitting involved in computing forward rates from Treasury coupon securities.

2.2 The Sensitivity of U.S. Interest Rates and Inflation Compensation to Economic News

The United States has issued inflation-indexed Treasury securities since 1997. A natural question arising from our estimates above, then, is to what extent the strong responses in far-ahead forward interest rates are due to changes in real interest rates, as opposed to changes

18. U.S. Treasury STRIPS (Separate Trading of Registered Interest and Principal Securities) are created by decoupling the individual coupon and principal payments from U.S. Treasury notes and bonds into pure zero-coupon securities. See Sack (2000) for more details and the potential usefulness of STRIPS for estimating the Treasury yield curve. In this paper, we compute the one-year forward rate ending in ten years using the nine-year STRIPS security and ten-year STRIPS security and applying equation (1).

in inflation compensation—the difference between nominal and real interest rates. Table 2 investigates this interesting question.

The primary shortcoming of U.S. Treasury inflation-indexed securities—commonly referred to as TIPS—is that they were issued for the first time in January 1997 and only annually for the first few years after that date. We therefore cannot compute a far-ahead forward real rate for the United States until January 1998, giving us a sample that covers only about seven and a half years. Nonetheless, the high frequency of the data still leaves us with almost a thousand observations with which to perform our analysis.

We obtained data on the forward real interest rates implied by TIPS from the Federal Reserve Board.[19] We define forward inflation compensation as the difference between the forward nominal rate and forward real rate at each horizon. This measure captures the compensation that investors demand both for expected inflation at the given horizon and for the risks or uncertainty associated with that inflation.[20]

In the first two columns of table 2, we repeat the regressions of the one-year spot rate and the ten-year-ahead one-year rate on our macroeconomic surprises over the sample of TIPS data (1998–2005). Our results over this sample are very similar to those in table 1, although the statistical significance is reduced for our coefficient estimates in both regressions. For example, only five of our twelve coefficients for the ten-year-ahead nominal rate are significant over this shorter sample, compared with nine of twelve in table 1, although the joint hypothesis that all coefficients are zero in that regression is still rejected at very high levels of statistical significance.[21] The signs and magnitudes of the coefficients in these two columns are also very similar to those we estimated over the larger 1994–2005 period.

19. The Federal Reserve Board provides real yield curve estimates beginning in January 1999. We extend the nine- to ten-year forward rate series back to January 1998 by taking the nine- and ten-year TIPS rates and computing the implied forward rate between the two using Shiller, Campbell, and Schoenholtz's (1983) approximation.

20. Forward real rates, nominal rates, and inflation compensation may also be affected by other factors, such as term premiums and premiums for liquidity. We discuss the robustness of all of our results with respect to these types of risk premiums in the next section.

21. The significance of the negative response of forward nominal rates to monetary policy surprises is notably absent over this later sample, perhaps reflecting the fact that these surprises become generally smaller and less frequent in the later part of our sample (Swanson, 2005). Another possible explanation for the smaller number of significant coefficients over the later sample is that long-term interest rates have gradually become better anchored in the United States. We leave this as an interesting question for future research.

Table 2. U.S. Forward Rate Responses to Domestic Economic News, 1998–2005[a]

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward real rate ending in ten years*	*One-year forward inflation compensation ending in ten years*
Capacity utilization	1.55***	0.91	0.51	0.40
	(2.92)	(1.33)	(1.31)	(0.66)
Consumer confidence	1.34**	0.50	0.18	0.32
	(2.57)	(0.75)	(0.47)	(0.55)
Core CPI	1.01	1.25	–0.37	1.63**
	(1.58)	(1.53)	(–0.80)	(2.28)
Employment cost index	1.14	1.13	–0.10	1.23
	(1.48)	(1.15)	(–0.17)	(1.43)
Real GDP (advance)	2.37***	1.91*	0.02	1.89**
	(2.92)	(1.84)	(0.04)	(2.08)
Initial jobless claims	–1.06***	–0.74**	–0.20	–0.54*
	(–4.25)	(–2.32)	(–1.09)	(–1.94)
NAPM/ISM manufacturing survey	2.26***	2.96***	1.74***	1.22**
	(4.39)	(4.49)	(4.59)	(2.12)
New home sales	0.23	0.67	–0.32	0.99*
	(0.51)	(1.15)	(–0.94)	(1.93)
Nonfarm payrolls	4.45***	1.79**	1.26***	0.54
	(8.02)	(2.52)	(3.07)	(0.88)

Table 2 (continued)

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward real rate ending in ten years*	*One-year forward inflation compensation ending in ten years*
Retail sales (excl. cars)	1.60***	1.52*	0.68	0.84
	(2.55)	(1.88)	(1.46)	(1.18)
Unemployment rate	–1.20*	0.89	0.84*	0.05
	(–1.95)	(1.13)	(1.85)	(0.07)
Monetary policy	0.36***	–0.01	0.01	–0.02
	(4.35)	(–0.13)	(0.18)	(–0.26)
No. 0bservations	950	950	950	950
R^2	0.15	0.06	0.04	0.04
Joint test p value	0.000***	0.000***	0.000***	0.010**

Source: Authors' computations.
* Statistically significant at the 10 percent level. ** Statistically significant at the 5 percent level. *** Statistically significant at the 1 percent level.
a. The sample is from January 1998 to October 2005, at daily frequency on the dates of macroeconomic and monetary policy announcements. Regressions also include a constant, a Y2K dummy that takes on the value of 1 on the first business day of 2000, and a year-end dummy that takes on the value of 1 on the first business day of any year (coefficients not reported). Macroeconomic data release surprises are normalized by their standard deviations, so these coefficients represent a basis point per standard deviation response. Monetary policy surprises are in basis points, so these coefficients represent a basis point per basis point response. Inflation compensation is the difference between nominal and real rates. Joint test p value is for the hypothesis that all coefficients (other than the constant and dummy variables) are zero. t-statistics are reported in parentheses.

In the third and fourth columns of table 2, we decompose the response of forward nominal rates into its constituent real rate and inflation compensation components. We find some evidence that part of the estimated responsiveness of nominal forward rates is actually due to movements in real interest rates, particularly for the NAPM/ISM manufacturing survey and nonfarm payrolls releases.[22] In the majority of cases, however, the responsiveness of long-term nominal interest rates is due at least partially to changes in inflation compensation. Five of our twelve estimated coefficients are statistically significant, and the joint hypothesis that all coefficients are zero is rejected with a *p* value of about 1 percent.

3. Possible Explanations for the Behavior of U.S. Long-Term Interest Rates

In steady state, the short-term nominal interest rate, i^*, equals the steady-state real interest rate, r^*, plus the steady-state level of inflation, π^*, by Fisher's equation:

$$i^* = r^* + \pi^* . \tag{5}$$

As mentioned above, standard asset-pricing theory indicates that forward rates with sufficiently long horizons—that is, $f_t^{(N)}$ for N large, where $f_t^{(N)}$ is the forward rate ending in N years' time—equal the expected steady-state short-term rate plus a risk premium, ρ:

$$f_t^{(N)} = r^* + \pi^* + \rho . \tag{6}$$

The fact that $f_t^{(N)}$ responds to many macroeconomic data releases and monetary policy surprises indicates that one (or more) of r^*, π^*, and ρ is changing in response to these surprises.

3.1 Some Nonexplanations for the Excess Sensitivity Puzzle: r^* and ρ

In our search for a solution to the excess sensitivity puzzle documented above, we consider, but ultimately discard, two possible causes: changes

22. We do not take a stand on why far-ahead real rates might move in response to economic news, although one possible explanation is that markets view the particular data release as informative about the economy's long-run rate of productivity growth and, hence, about the equilibrium real interest rate.

in r^* (the long-run equilibrium real interest rate) and changes in ρ (the risk premium). Although r^* is a potentially time-varying component of steady-state short-term rates, our results for the nominal forward rates are probably not due to r^* responding to surprises. We have two reasons for ruling out time variance in steady-state real rates as the main culprit. First, TIPS provide a measure of far-ahead forward real rates, and as we showed in table 2, the sensitivity of nominal rates in the United States to economic news was almost always attributable to changes in inflation compensation, rather than to changes in real rates. Second, many of the nominal interest rate responses that we estimate are difficult to interpret in terms of changes in r^*. For example, it is difficult to explain why a surprise uptick in inflation (of either the CPI or the PPI) would lead the market to revise upward its estimate of r^*, the long-run equilibrium real rate of interest.[23] Similarly, a surprise monetary policy tightening is not likely to lead the market to revise its estimate of r^* downward—presumably, a surprise tightening of policy, to the extent that it provides any information about r^*, indicates that the FOMC views r^* as being higher than the market estimate.

This is not to say that changes in the market's perception of r^* are necessarily unimportant. Indeed, changes in r^* may have had some effect on long-term interest rates in our sample, particularly in the late 1990s, when market estimates of the long-run rate of productivity growth in the United States were largely in flux. Relying solely on changes in r^* to explain our empirical results, however, is likely to cause difficulties for precisely the reasons described above.

Alternatively, one might argue that changes in the risk premium, ρ, are the most likely explanation for our findings of excess sensitivity in long-term interest rates. While some authors find little evidence for time-varying risk premiums in the data (for example, Bekaert, Hodrick, and Marshall, 2001), a number of prominent studies (such as Fama and Bliss, 1987; Campbell and Shiller, 1991) document strong violations of the expectations hypothesis for a wide variety of samples and securities, suggesting that the risk premiums embedded in long-term bond yields may, in fact, vary substantially over time. A time-varying risk premium is often offered as an explanation for the excess volatility puzzle and as a likely factor in the failure of the expectations hypothesis for longer maturities.

23. Even if one regards surprises in inflation as being informative about productivity growth in the late 1990s, the usual story that is told is that surprisingly low inflation was indicative of high productivity growth, which would, in turn, be related to a higher equilibrium real rate, r^*.

For our analysis, however, as long as the variation in risk premiums is small enough at the very high frequencies we consider, the change in bond yields over the course of the day will effectively difference out the risk premium at each point in our sample, allowing us to interpret the change in yields as being driven primarily by the change in expectations. While there is no a priori reason why risk premiums should vary only at lower frequencies, the predictors of excess returns on bonds emphasized in the studies above generally have this feature—that is, the variation from one day to the next is very small, while the large variations in premiums that they estimate occur at much lower frequencies, particularly the business cycle (Cochrane and Piazzesi, 2005; Piazzesi and Swanson, 2004). Thus, the failure of the expectations hypothesis alone is not sufficient to call our analysis into question.

Nevertheless, risk premiums are poorly understood, so the fact that previous estimates of time-varying risk have generally found predictability only at lower frequencies does not imply that they could not change appreciably from one day to the next. In order for changes in risk premiums to explain our results, however, one would have to explain why they would move so systematically in the way that we document, being positively correlated with output and inflation news while moving inversely with surprises in monetary policy.[24] Moreover, one would have to explain why we do not find similar movements in risk premiums in the United Kingdom or Sweden, as documented in Gürkaynak, Levin, and Swanson (2006)—if anything, one would expect the importance of risk premiums to be greater in these smaller, less liquid markets—or why the behavior of risk premiums in the United Kingdom would have changed after the Bank of England gained independence from Parliament in 1997 (Gürkaynak, Sack, and Swanson, 2003; Gürkaynak, Levin, and Swanson, 2006).

Given that current theory puts little structure on the behavior of term premiums, one can write an ad hoc model of the term premium that would match our empirical findings. However, the fact that we did not observe a strong response of real interest rates to economic news

24. Cochrane and Piazzesi (2005) and Piazzesi and Swanson (2004) find that risk premiums in Treasury securities and interest rate futures move countercyclically over the business cycle. This is exactly opposite to the direction that would be needed to explain our findings that far-ahead forward interest rates in the United States and in the United Kingdom before central bank independence comove positively with surprises in output and employment.

in the United States suggests that if changes in risk premiums are responsible for the excess sensitivity of the forward nominal rates, any such risk seems to be more closely related to inflation compensation than to real rates. This is in line with our interpretation that it is the perceived *distribution* of future inflation outcomes (and not necessarily only its mean) that is unanchored.

3.2 A Possible Explanation for Excess Sensitivity: Changes in π^*

While we do not wish to discount the importance of changes in market perceptions of r^* or changes in risk premiums that are unrelated to inflation, we find each of them inadequate on its own to explain all of our empirical results. We now show that changes in the market's perception of π^*, the long-run inflation objective of the central bank, helps explain all of our findings. Thus, changes in π^* are not only necessary for explaining at least some of our results, but also sufficient.[25]

Model with time-varying π^* and perfect information

We demonstrate the sufficiency of changes in π^* by augmenting the benchmark model from section 1 to include an additional equation that permits the central bank's inflation objective to vary over time, without taking a stand on why this might be so. In this alternative specification, past values of inflation affect the central bank's inflation target. Our assumed functional form for the time-variance in π^* is

$$\pi_t^* = \pi_{t-1}^* + \theta(\overline{\pi}_{t-1} - \pi_{t-1}^*) + \varepsilon_t^{\pi^*}, \tag{7}$$

where $\overline{\pi}_{t-1}$ is the trailing four-quarter moving average of inflation. Thus, persistently low (high) inflation will, over time, tend to decrease (increase) the central bank's long-run inflation target.[26] Exogenous changes in the central bank's inflation objective, π^*, are captured by the shock $\varepsilon_t^{\pi^*}$.

25. While the model presented below is based on time variance in the perceived mean of the steady-state inflation distribution, the results would go through if other moments of that distribution were time varying, as well. These would be reflected in the inflation term premium.

26. This has some similarities to the idea of opportunistic disinflation described in Orphanides and Wilcox (2002).

Our benchmark model with time-varying π^* thus takes the form:

$$\pi_t = \mu_\pi E_t \pi_{t+1} + (1-\mu_\pi) A_\pi(L)\pi_t + \gamma y_t + \varepsilon_t^\pi, \tag{8}$$

$$y_t = \mu_y E_t y_{t+1} + (1-\mu_y) A_y(L) y_t - \beta(i_t - E_t \pi_{t+1}) + \varepsilon_t^y, \tag{9}$$

$$i_t = (1-c)\left[\bar{\pi}_t + a(\bar{\pi}_t - \pi_t^*) + b y_t\right] + c\, i_{t-1} + \varepsilon_t^i, \text{ and} \tag{10}$$

$$\pi_t^* = \pi_{t-1}^* + \theta(\bar{\pi}_{t-1} - \pi_{t-1}^*) + \varepsilon_t^{\pi^*}, \tag{11}$$

where equation (10) now explicitly recognizes the existence of a time-varying inflation target. We use the same parameter values for the model as for the Rudebusch specification in section 1, and we select a value for θ to roughly calibrate our impulse response functions to match the estimated responsiveness of long-term forward rates in our data. It turns out that we require relatively small values for θ (the loading of the central bank's inflation target on the past year's inflation) to match the term structure evidence. We thus set θ equal to 0.02 for the simulations below, implying that annual inflation one percentage point above target leads the central bank to raise its target by 2 basis points. This may seem negligibly small, but the persistence of inflation—particularly the four-quarter trailing average that enters into equation (11)—leads to cumulative effects on π^* that are nonnegligible, as we now show.

Figure 2 plots the impulse responses of inflation, the output gap, the short-term interest rate, and π^* to a one percent shock to each of equations (8) through (11).[27] The qualitative features of our empirical findings are reproduced very nicely. For example, after a one percent inflation shock (the first column), the short-term nominal interest rate rises gradually, peaks after a few years, and then returns to a long-run steady-state level that is about 35 basis points higher than the original steady state. This is due to the fact that the higher levels of inflation on the transition path cause the central bank's long-run objective, π^*, to rise. A similar response of short-term nominal interest rates and inflation can be seen in response to a one percent shock

27. The model has no indexation to steady-state inflation, so the central bank's π^* does not enter the private sector's equations directly. Rather, it only enters indirectly through the private sector's forecast of π_{t+1} and y_{t+1}, which depend on the current and expected future path for the interest rate (which depends on π^*).

to output (the second column). For the federal funds rate shock (the third column), as inflation in the economy falls in response to the monetary tightening, the central bank's long-run target π^* gradually falls, as well. In the long run, the short-term nominal interest rate and inflation settle below their initial levels, producing exactly the kind of inverse relation between far-ahead forward rates and short rates that we found in the data.

Model with time-varying π^* and imperfect information

The above model can also be extended to include the case in which the private sector does not directly observe the central bank's inflation objective, π^*, and thus must infer it from the central bank's actions, as in Kozicki and Tinsley (2001), Ellingsen and Soderstrom (2001), and Erceg and Levin (2003). The advantages of a model with imperfect information are threefold. First, it emphasizes that the behavior of the term structure is driven by private sector expectations of future outcomes, which in the case of imperfect information can differ from the actual impulse responses to a particular (unobserved or imperfectly observed) shock. Second, a model with imperfect information provides a more realistic description of long-term interest rate behavior in the United States, since the Federal Reserve's long-term objective for inflation, π^*, is unknown to financial markets. Third, the presence of imperfect information increases the importance and effects of monetary policy shocks in the model, which allows for a better calibration to our empirical results than the model with perfect information can provide.

To consider the case of imperfect information, equations (8) through (11) must be augmented to include a private sector Kalman filtering equation:

$$\hat{\pi}_t^* = \hat{\pi}_{t-1}^* + \theta\left(\overline{\pi}_{t-1} - \hat{\pi}_{t-1}^*\right) - \kappa\left(i_t - \hat{i}_t\right). \tag{12}$$

For simplicity and tractability, we assume that the forms of equations (8) through (11), all parameter values, and the shocks ε^π and ε^y are perfectly observed by the private sector. Thus, only π^*, ε^{π^*}, and ε^i are unobserved. Private agents update their estimate of the central bank's inflation target, denoted $\hat{\pi}_t^*$, using equation (12).[28] In

28. This procedure is optimal under the assumptions of normally distributed shocks and a normally distributed prior for the inflation target. For other shock distributions, the Kalman filter is the optimal linear inference procedure.

Figure 2. Impulse Responses with Time-Varying π^* (Perfect Information)

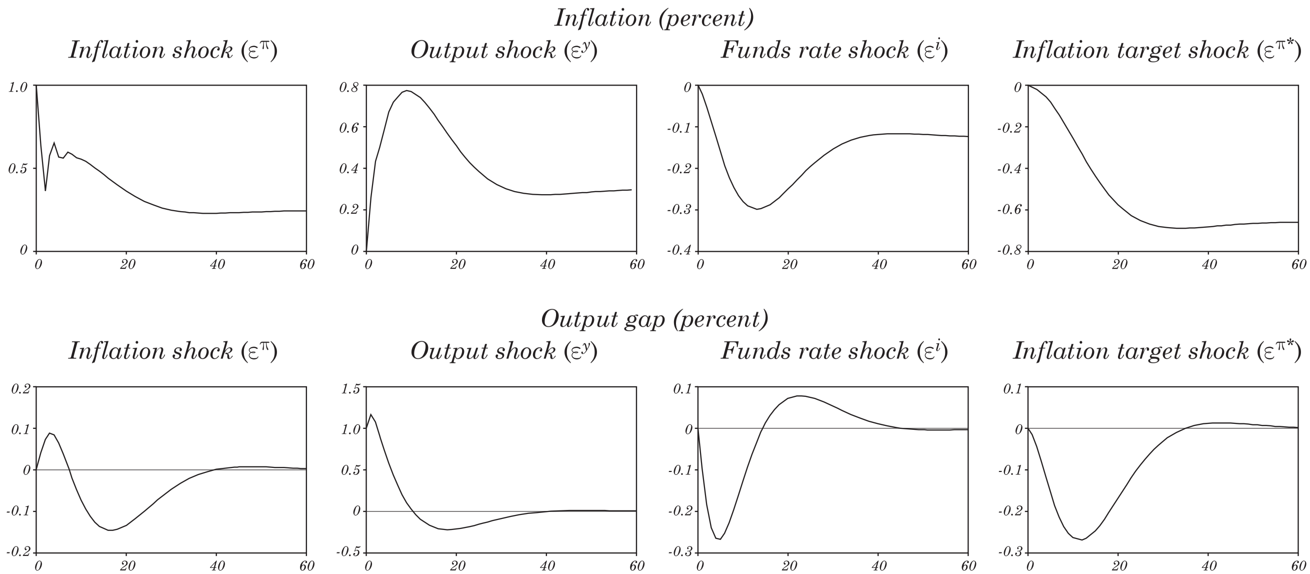
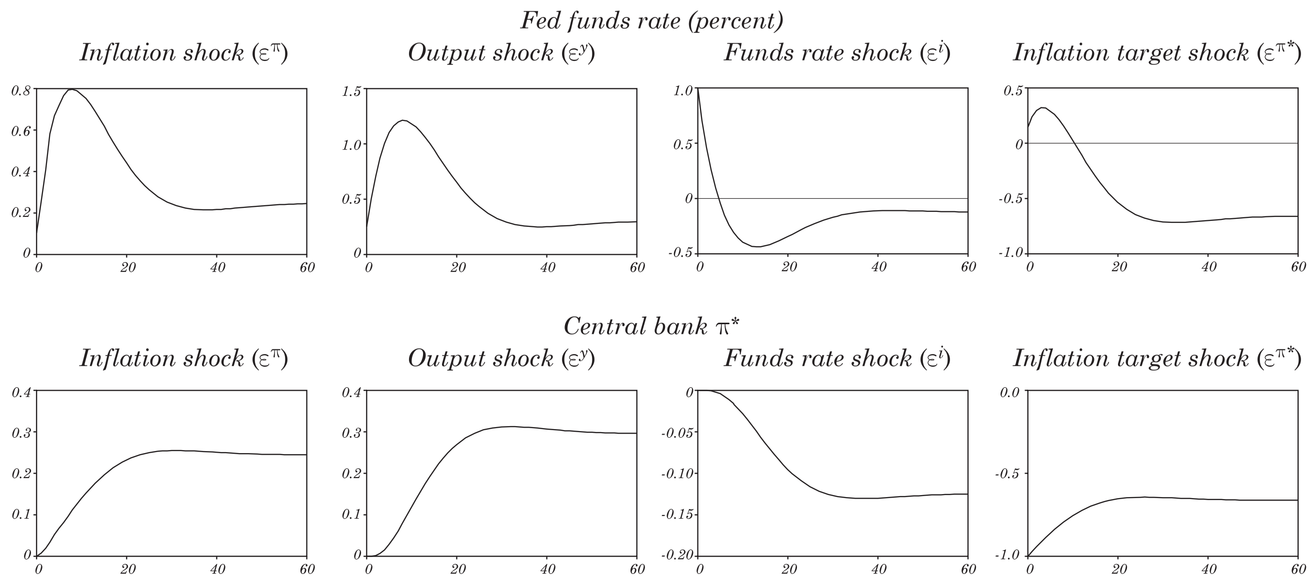


Figure 2 (continued)



Source: Authors' computations.

particular, agents observe the deviation of the interest rate from their expectation, $i_t - \hat{i}_t$, where $\hat{i}_t$ is obtained by substituting $\hat{\pi}^*_t = \hat{\pi}^*_{t-1}$ and $\varepsilon^i_t = 0$ into equation (10), and they revise $\hat{\pi}^*_t$ by an amount determined by the Kalman gain parameter, κ. Again, we choose (rather than estimate) a value for κ of 0.1, which is meant to be illustrative and matches the data.[29]

Figure 3 presents the private sector's expected impulse responses to inflation, the output gap, the short-term interest rate, and the central bank's inflation objective following a shock to each of equations (8) through (11). Because this version of the model features imperfect information, the impulse responses expected by the private sector on impact may differ from the actual impulse responses from a shock. In particular, the private sector is initially unable to distinguish between the temporary shock, ε^i, and the permanent central bank preference shock, ε^{π^*}. The expected impulse responses to those two shocks are therefore identical, up to a scale factor, even though the actual impulse responses to those two shocks play out quite differently over time.[30]

29. Alternatively, one could derive the optimal value for κ from the variance of the shocks to π^* and to *i,* but this value would have to be indirectly inferred anyway since π^* is unobserved. The value of 0.1 that we use for κ corresponds to a ratio of $\sigma_i/\sigma_{\pi^*} = 3$.

30. Expected and actual impulse responses for the case of imperfect information are calculated as follows. If, starting from steady state, the model is hit by a shock to π or to y, then the private sector observes those two shocks, so there is no imperfect information and the impulse responses are just like in the perfect information case. If, instead, there is a shock to i or to the central bank's π^*, then the private sector does not observe the true shock and must estimate what the shock was from the observed change in i. The private sector optimally assigns part of the change in i to ε^i and part of the change in i to ε^{π^*}. Knowing the true structure of the economy, the private sector then projects the economy forward using its above two estimates for the shocks to i and to π^*. This yields the *expected* impulse response functions at time t. This solution also yields the *actual* equilibrium of the model at time t (and time t only). In period t+1, the economy will evolve slightly differently than the private sector had expected the previous period (because the private sector did not observe the true shocks to i and π^*). In particular, i will be a little different again from what the private sector was expecting, so agents will think that their previous estimate of π^* may have been wrong or that there may have been another shock to i or another shock to π^*. (Of course, in an impulse response function, we do not hit the model with any additional shocks, but the private sector does not know this). The private sector thus optimally updates its estimate of π^* again, and projects the economy forward again using the true structure. This solution yields the equilibrium of the model at time $t + 1$ (and time $t + 1$ only). Come period $t + 2$, the economy will evolve slightly differently than the private sector had expected the previous period, and so forth. We repeat this procedure to obtain the entire actual response of interest rates to the shock (which we plot in figure 4). Again, the private sector's estimate of π^* does not enter the private sector's equations directly, but only indirectly through the private sector's forecast of π_{t+1} and y_{t+1}, which depends on the current and expected future path of the interest rate, which in turn depends on the private sector's estimate of π^*.

The expected impulse responses in figure 3 again reproduce the qualitative features of our empirical findings nicely. The responses to an inflation shock (the first column) or an output shock (second column) are identical to the perfect information case in figure 2, because we have assumed that the private sector has perfect information regarding those two variables. For the case of the federal funds rate shock (third column), however, two effects are now present. First, when the private sector sees the surprise tightening in short-term interest rates, they cannot tell whether the shock is purely temporary (ε^i) or reflects a more permanent change in π^*, so they respond to the shock by partially revising downward their estimate of the central bank's π^*. Inflation in the economy thus falls in response to both the monetary tightening and the fall in inflation expectations, leading to larger effects than in the perfect information case. Second, the central bank's long-run objective, π^*, falls over time as inflation comes in below target, as was true in the perfect information case. The effect of the additional channel arising from imperfect information is to increase the relative size and importance of the effects of the interest rate shock on the term structure, allowing for a better calibration to our empirical results and providing a more realistic model of long-term interest rates in the United States.

Note that imperfect information about the central bank's target, π^*, plays a role only in the third and fourth columns of the figure. A model based solely on imperfect information or imperfect credibility, as in Kozicki and Tinsley (2001) or Erceg and Levin (2003), would be unable to reproduce our findings of excess sensitivity of U.S. interest rates to output and inflation surprises as long as shocks to ε^π and ε^y are observed.

For reference, the actual impulse responses of the model (equations 7 through 12) are depicted in figure 4. The figure illustrates how the differing effects of shocks to i and to π^* play out over time. The fifth row depicts the evolution of the private sector's estimate, $\hat{\pi}^*$, in response to each shock. Shocks to inflation or output, about which there is no imperfect information, lead to responses of $\hat{\pi}^*$ that are identical to those of π^*, but the two variables evolve differently for the imperfectly observed cases of shocks to i and π^*.

Finally, our hypothesis that the private sector's expectations of the central bank's long-run inflation objective, π^*, have varied over time is also consistent with measures of these expectations derived from survey data. For example, the median ten-year CPI inflation forecast

Figure 3. Expected Impulse Responses with Time-Varying π^* (Imperfect Information)

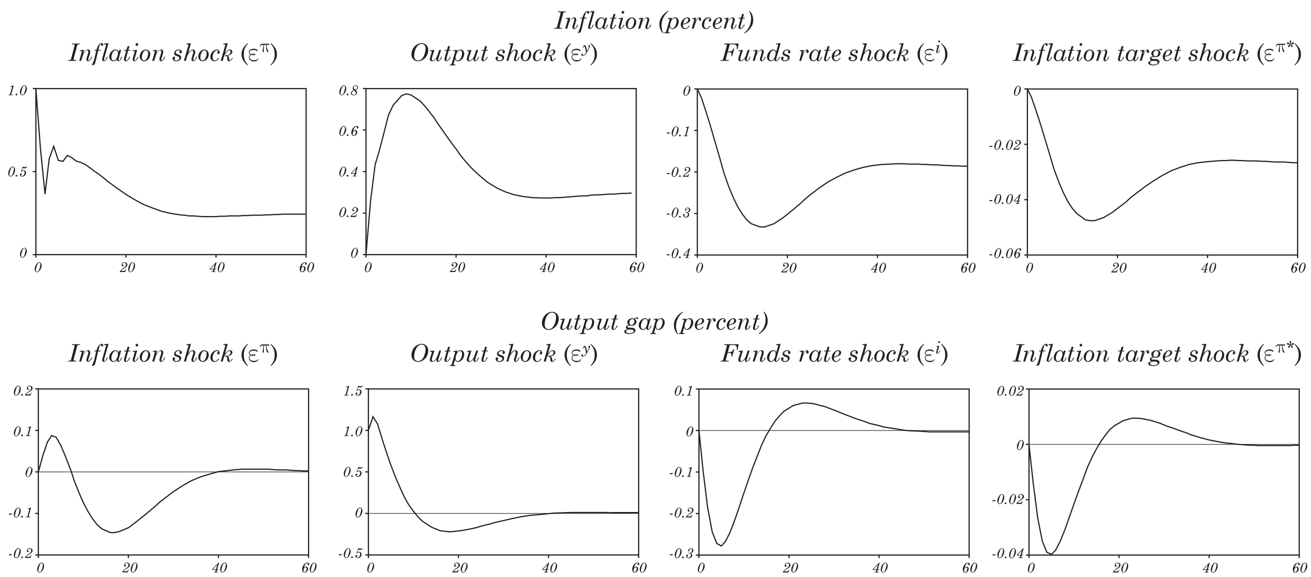
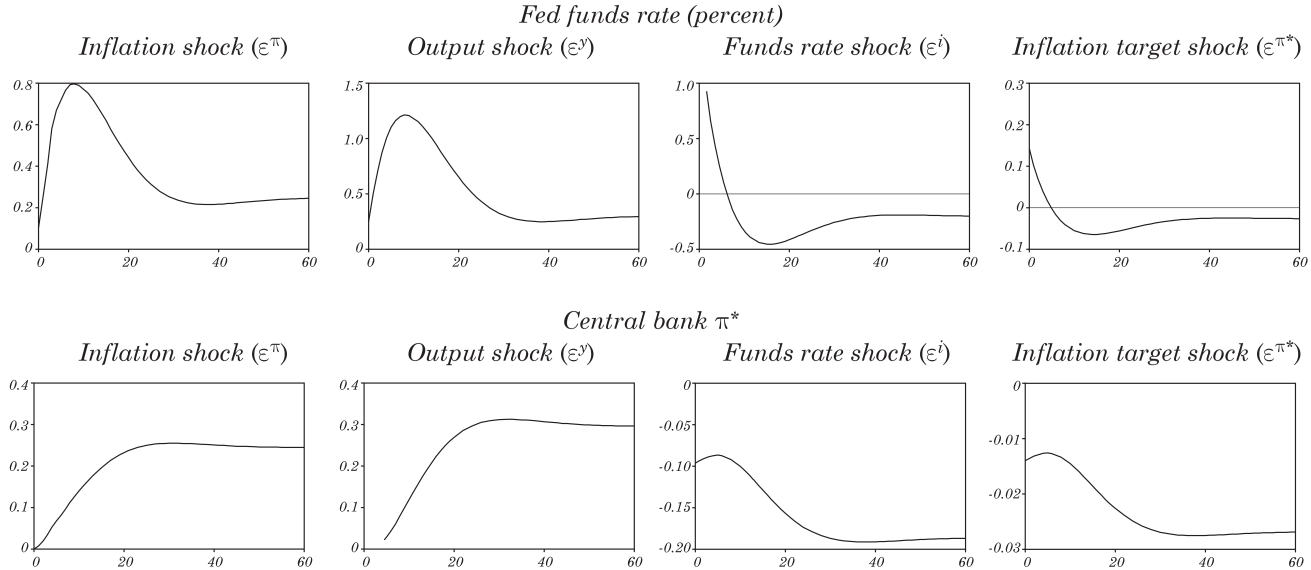


Figure 3 (continued)



Source: Authors' computations.

Figure 4. Actual Impulse Responses with Time-Varying π^* (Imperfect Information)

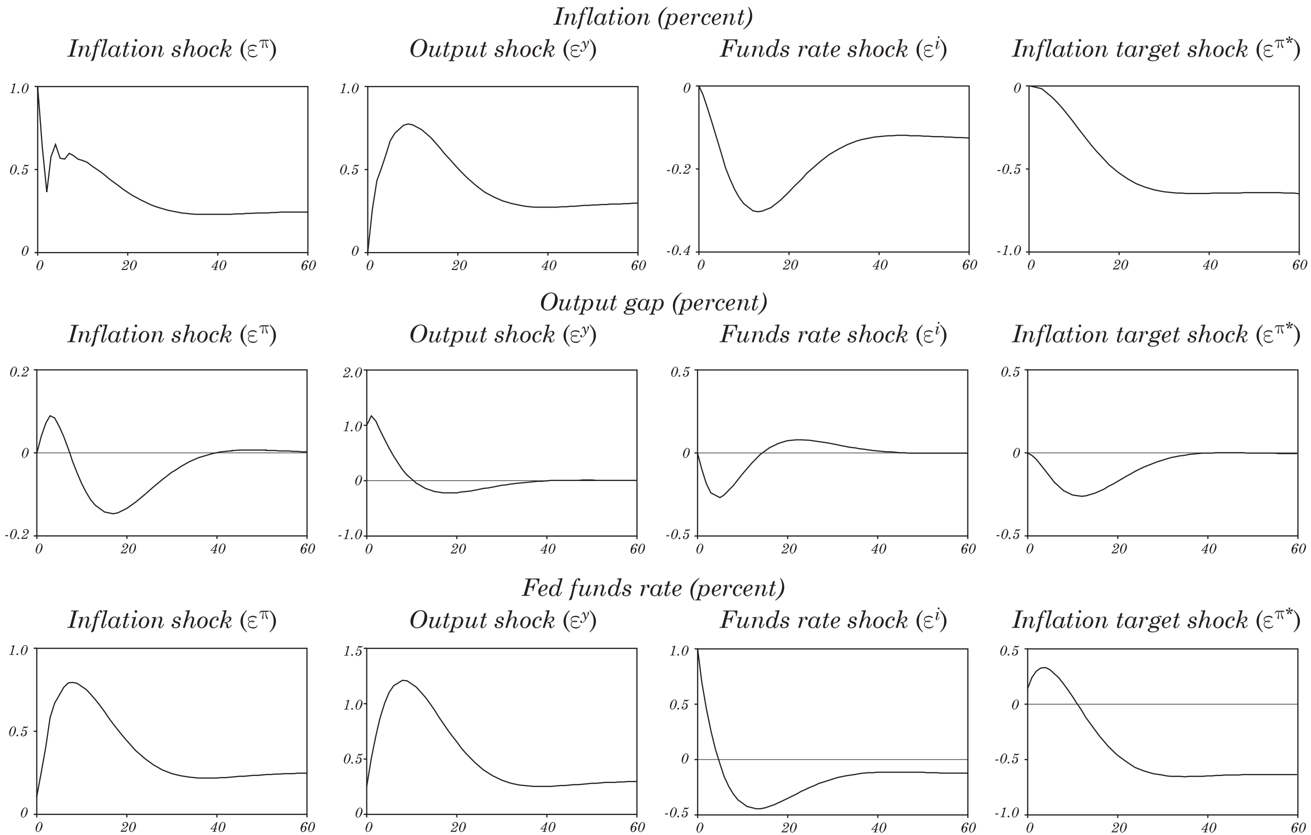
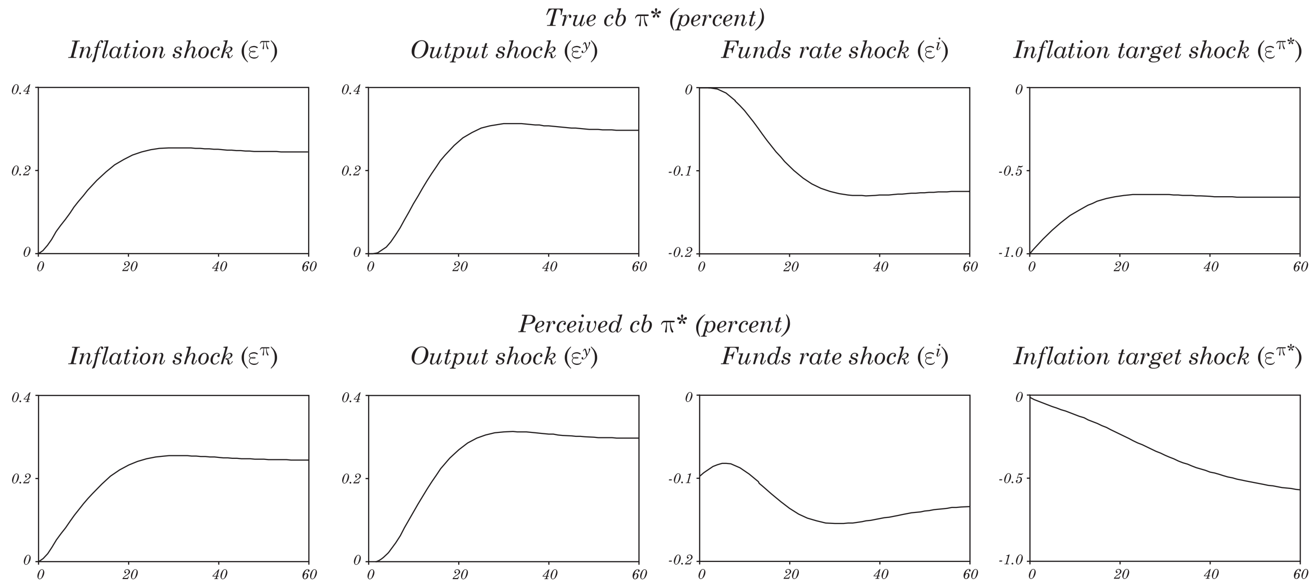


Figure 4 (continued)



Source: Authors' computations.

in the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters fell from 4 percent in the fourth quarter of 1991 (the first time the long-run forecast question was asked) to a little under 2.5 percent by the end of 2002. This decline of about 1.5 percentage points compares with a fall of about 2.5 percentage points in ten-year nominal forward interest rates over the same period.

4. The Sensitivity of Long-Term Interest Rates in Canada and Chile

We have shown that U.S. long-term interest rates are excessively responsive to economic news, and that this responsiveness is well explained by changes in financial market perceptions of a long-run inflation objective in the United States that is not well anchored. We now explore whether long-term interest rates are any more stable in countries that are explicit inflation targeters than in the United States. Gürkaynak, Levin, and Swanson (2006) consider the cases of Sweden and the United Kingdom and find that far-ahead forward interest rates are much better anchored in those two countries than in the United States. In this paper, we extend the comparison to Canada and Chile, which have been formal inflation targeters throughout much of the 1990s and 2000s.[31] Despite these relatively short sample periods, our high-frequency methodology provides us with several hundred to a thousand observations for each of these countries for our analysis.

4.1 The Sensitivity of Long-Term Interest Rates in Canada

We obtained data on Canadian macroeconomic news releases and financial market expectations of those releases from two sources:

31. Both Canada and Chile adopted an inflation-targeting framework in which the target was not firmly anchored at first, but was rather successively lowered during a transition period. Canada adopted its inflation-targeting framework in 1991, but the target was not stabilized at the current level of 1–3 percent until early 1995. Chile, in turn, adopted its inflation-targeting framework in 1991, but the target was not stabilized at the current level of 2–4 percent until early 2001. For our purposes, the latter dates are the more relevant ones. Finally, the adoption of an inflation-targeting range rather than a point makes very little difference in theory, because the optimal monetary policy is always to aim for inflation to lie at the midpoint of the range, as discussed, for example, by Orphanides and Wieland (2000).

Money Market Services and Bloomberg.[32] When those data sets overlap, they agree very closely. Between these two sources, we have data on Canadian capacity utilization, the consumer price index, core consumer price index, employment, real GDP, retail sales, the unemployment rate, and wholesale trade. Most of these series go back to 1996, and a few go back even farther.[33] To measure the surprise component of Canadian monetary policy announcements, we obtained the dates of changes in the Bank of Canada's target overnight interbank rate back to 1995 from the Bank of Canada's web site, and we measured the surprise component of these changes as the change in the three-month Canadian Treasury bill on the dates of these monetary policy changes.

We obtained data on Canadian nominal bond yields from the Bank of Canada's web site and data on real bond yields from Bloomberg. The Bank of Canada provides nominal zero-coupon yield curve estimates extending back to the 1980s. Inflation-indexed bond data for Canada is more limited: Canada issued its first inflation-indexed bond in 1991 and its second in 1996, implying that we cannot compute a forward real rate for Canada until 1996. Moreover, Canada has issued indexed bonds only at the thirty-year maturity. These securities thus have extremely long durations and appeal primarily to pension funds, insurance companies, and individual investors, resulting in low levels of secondary market liquidity, high transactions costs, and observed real interest rates that are noisy, particularly in the earlier years of our sample.[34] Thus, to reduce the noisiness of the data and facilitate

32. Data from Bloomberg were freely available to us through a subscription at the Federal Reserve Board and the Federal Reserve Bank of San Francisco. However, Money Market Services (our source for the U.S. data) had data for a number of Canadian series that were not covered by Bloomberg and that we thought might be important, so we purchased these additional series from Money Market Services. See the appendix for details.

33. Details of the data are provided in the appendix.

34. To compute far-ahead forward real rates in Canada, we use as many of the 2021, 2026, 2031, and 2036 maturity coupon bond yields as are available on any given date and compute the far-ahead forward rates between pairs of securities using Shiller, Campbell, and Schoenholtz's (1983) approximation. We use the average one-day change in these forward rates in our regressions. We use a longer (twenty- to thirty-year-ahead) horizon to proxy for the nine-year-ahead real one-year forward rate in Canada, because we simply do not have nine-year-ahead Canadian indexed bond data. Although we could use a twenty- or thirty-year-ahead horizon for our nominal Canadian forward rates, as well, we judged that the lower liquidity and higher trading costs of these longer-horizon securities would more than offset any gains from having a precise match in maturity.

comparison with the United States, we begin our analysis of Canada in January 1998.[35]

The results of our analysis for Canada are presented in tables 3 and 4. Table 3 investigates the sensitivity of Canadian far-ahead forward interest rates and inflation compensation to domestic economic news. As in previous tables for the United States, the first column reports the response of the one-year Canadian nominal spot rate to domestic news releases. Short rates respond significantly to several of the statistics we consider, with signs and magnitudes that are consistent with our earlier estimates for the United States. In sharp contrast to the United States, however, far-ahead forward nominal rates in Canada (in the second column) respond significantly to almost none of these news releases: only the coefficient on monetary policy surprises is significant at even the 10 percent level, and that result is driven by just one or two observations at the beginning of the sample. We find very similar results when we look at far-ahead forward inflation compensation (the fourth column). Here again, only one coefficient is marginally statistically significant (on the core CPI), and even that coefficient seems to be driven by a puzzling negative relation between far-ahead forward real interest rates and core CPI releases. The joint hypothesis that all coefficients in the regression are equal to zero in these two regressions is not rejected at any standard level of significance.

In table 4, we explore whether Canadian far-ahead forward interest rates and inflation compensation respond to U.S. economic data releases and monetary policy announcements. Because Canada is a relatively small open economy, it is reasonable to think that short-term interest rates and even long-term real rates in Canada might be largely determined by developments in the rest of the world, particularly developments in the United States. We would still expect the long-run values of purely nominal variables, such as inflation and inflation expectations, to be determined primarily by domestic monetary policy, particularly at the far-ahead horizons

35. In 1996 and 1997, there are seven forward real rate changes in Canada of 100 basis points or more in a single day, and seventeen changes of 50 basis points or more. We believe that these observations are due to low trading volumes and low liquidity for these securities, rather than to perceived changes in economic fundamentals. After January 1998, there are no changes of 50 basis points or more. While noise and low liquidity may still be an issue in the indexed bond data after January 1998, we found that problems related to regression outliers were essentially eliminated by restricting attention to the post-1997 period. Moreover, this period matches our sample for the United States, allowing for closer comparability between our U.S. and Canadian results.

Table 3. Canadian Forward Rate Responses to Domestic Macroeconomic News, 1998–2005[a]

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward real rate ending in ten years*	*One-year forward inflation compensation ending in ten years*
Capacity utilization	0.19	0.61	0.59	0.02
	(0.16)	(0.39)	(0.85)	(0.01)
CPI	1.49*	−0.27	−0.79	0.53
	(1.68)	(−0.24)	(−1.61)	(0.47)
Core CPI	1.22	−0.23	−1.07**	0.84
	(1.58)	(−0.23)	(−2.49)	(0.86)
Employment	3.07***	0.65	0.34	0.31
	(4.48)	(0.75)	(0.90)	(0.35)
Real GDP	−1.01	−2.35	0.25	−2.60
	(−0.58)	(−1.08)	(0.26)	(−1.19)
Retail sales	1.48**	−0.29	0.00	−0.30
	(2.28)	(−0.36)	(0.01)	(−0.36)
Unemployment rate	0.31	−0.29	0.11	−0.40
	(0.50)	(−0.37)	(0.32)	(−0.51)
Wholesale trade	0.09	−0.55	−0.28	−0.27
	(0.13)	(−0.59)	(−0.69)	(−0.29)
Monetary policy	0.81***	−0.28	−0.06	−0.22
	(5.77)	(−1.57)	(−0.761)	(−1.25)
No. observations	327	327	327	327
R^2	0.19	0.02	0.10	0.03
Joint test p value	0.000***	0.806	0.006***	0.732

Source: Authors' computations.

* Statistically significant at the 10 percent level. ** Statistically significant at the 5 percent level. *** Statistically significant at the 1 percent level.

a. The sample is from January 1998 to October 2005, at daily frequency on the dates of macroeconomic and monetary policy announcements. Regressions also include a constant, a Y2K dummy that takes on the value of 1 on the first business day of 2000, and a year-end dummy that takes on the value of 1 on the first business day of any year (coefficients not reported). Macroeconomic data release surprises are normalized by their standard deviations, so these coefficients represent a basis point per standard deviation response. Monetary policy surprises are in basis points, so these coefficients represent a basis point per basis point response. Inflation compensation is the difference between nominal and real rates. Joint test p value is for the hypothesis that all coefficients (other than the constant and dummy variables) are zero. t-statistics are reported in parentheses.

Table 4. Canadian Forward Rate Responses to U.S. Macroeconomic News, 1998–2005[a]

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward real rate ending in ten years*	*One-year forward inflation compensation ending in ten years*
U.S. capacity utilization	1.42**	0.72	0.12	0.60
	(2.13)	(0.81)	(0.26)	(0.63)
U.S. consumer confidence	1.35*	–0.00	0.62	–0.62
	(1.91)	(–0.00)	(1.32)	(–0.61)
U.S. core CPI	0.96	2.07**	–0.30	2.37**
	(1.22)	(1.98)	(–0.59)	(2.10)
U.S. employment cost index	1.11	2.09	0.62	1.47
	(1.13)	(1.60)	(0.96)	(1.04)
U.S. real GDP (advance)	2.40**	0.40	–0.06	0.46
	(2.49)	(0.32)	(–0.09)	(0.34)
U.S. initial jobless claims	–0.99***	–0.72	–0.27	–0.45
	(–2.85)	(–1.56)	(–1.20)	(–0.89)
U.S. NAPM/ISM manufacturing survey	1.72**	1.88*	1.18**	0.69
	(2.18)	(1.79)	(2.27)	(0.61)
U.S. new home sales	–0.66	0.60	–0.52	1.12
	(–1.22)	(0.85)	(–1.47)	(1.46)
U.S. nonfarm payrolls	4.32***	1.66*	1.78***	–0.13
	(6.63)	(1.92)	(4.16)	(–0.14)
U.S. retail sales (excl. cars)	1.12	0.47	0.18	0.29
	(1.39)	(0.44)	(0.35)	(0.25)

Table 4 (continued)

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward real rate ending in ten years*	*One-year forward inflation compensation ending in ten years*
U.S. unemployment rate	–1.04 (–1.31)	–1.72 (–1.63)	0.42 (0.80)	–2.13* (–1.87)
U.S. monetary policy	0.37*** (3.52)	–0.20 (–1.45)	0.14** (2.03)	–0.34** (–2.27)
No. observations	939	939	939	939
R^2	0.16	0.03	0.06	0.02
Joint test p value	0.000***	0.148	0.001***	0.361

Source: Authors' computations.
* Statistically significant at the 10 percent level. ** Statistically significant at the 5 percent level. *** Statistically significant at the 1 percent level.
a. The sample is from January 1998 to October 2005, at daily frequency on the dates of macroeconomic and monetary policy announcements. Regressions also include a constant, a Y2K dummy that takes on the value of 1 on the first business day of 2000, and a year-end dummy that takes on the value of 1 on the first business day of any year (coefficients not reported), and Canadian macroeconomic news releases (coefficients not reported since they are very similar to table 3). Macroeconomic data release surprises are normalized by their standard deviations, so these coefficients represent a basis point per standard deviation response. Monetary policy surprises are in basis points, so these coefficients represent a basis point per basis point response. Inflation compensation is the difference between nominal and real rates. Joint test p value is for the hypothesis that all coefficients (other than the constant and dummy variables) are zero. t-statistics are reported in parentheses.

we are considering in this paper. Thus, while short-term rates and perhaps long-term real interest rates in Canada might be expected to respond to U.S. economic news, we would still expect far-ahead forward inflation compensation and perhaps nominal rates to remain largely invariant, if financial markets view the distribution of long-run inflation outcomes in Canada as being well-anchored.

The regressions in table 4 include both Canadian and U.S. macroeconomic data releases and monetary policy announcements, although coefficients on the Canadian releases are not reported to save space (they are very similar to those reported in table 3). The first column of table 4 shows that short-term interest rates in Canada are indeed significantly affected by U.S. monetary policy announcements and by many U.S. macroeconomic data releases. Still, far-ahead forward nominal rates in the second column) are not very responsive to these U.S. economic news releases, with three coefficients exhibiting only a marginal degree of statistical significance. The joint hypothesis that all coefficients are zero in this far-ahead forward nominal rate regression is not rejected at any standard level of statistical significance. The same observations generally remain true when we look at far-ahead forward inflation compensation (the fourth column): although this period includes three U.S. data releases that are significantly related to Canadian far-ahead forward inflation compensation at the 10 percent level or better, the joint test that all coefficients are equal to zero is not rejected at any standard level of significance.

These findings for Canada are reminiscent of those reported by Gürkaynak, Levin, and Swanson (2006) for the United Kingdom and Sweden, which were both inflation targeters over much of the 1990s. In their analysis, the United Kingdom and Sweden displayed a much greater anchoring of far-ahead forward nominal rates and inflation compensation in response to economic news than did the United States. Finally, in the case of the United Kingdom, the Bank of England was granted operational independence from Parliament in 1998. Gürkaynak, Levin, and Swanson show that, while the United Kingdom has had substantially better-anchored long-term inflation expectations than the United States since that date, the data for the early 1990s display a sensitivity of forward nominal rates and inflation compensation that is very similar to what we observe in the United States. All of these findings support the conclusion that a credible inflation-targeting framework significantly helps to anchor the private sector's perception of the distribution of future long-run inflation outcomes.

4.2 The Sensitivity of Long-Term Interest Rates in Chile

Chile has a much less extensive set of monthly macroeconomic data releases than are available in a more industrialized country such as the United States or Canada. We obtained data on Chilean monthly macroeconomic data releases and ex ante private sector forecasts of these releases from the Central Bank of Chile for four macroeconomic statistics: consumer price index inflation, monetary policy announcements, real GDP growth in the current quarter, and real GDP growth in the previous quarter. However, whereas our forecast data for the United States and Canada are at most a few days old on release, the Chilean data can be as much as two or even three weeks old by the time of the actual release, because the private sector macroeconomic forecast is only collected every few weeks. Thus, our measure of macroeconomic surprises for Chile is likely to suffer from measurement error, which will diminish our chances of finding statistically significant effects of releases on interest rates at even the short end of the yield curve.[36]

The Central Bank of Chile also provided us with Chilean real and nominal yield curve data. In contrast to the United States and Canada, there were no long-term nominal government bonds outstanding in Chile until 2002—all long-term government debt issued prior to that date was inflation indexed, at least in the last thirty years. This lack of long-term nominal debt presumably reflects the fact that the Chilean government was unwilling to pay the large risk premiums that investors would have demanded to hold such long-term nominal liabilities during a period in which markets viewed the government and the central bank as being greater credit and inflation risks than they are today. Thus, our sample for Chile is restricted to the 2002–05 period, which, although very short, still provides us with about four hundred observations for our analysis given the high frequency of the data. Moreover, even with ideal data, it would be difficult to extend our sample for Chile further back than 2001: although Chile formally adopted an inflation-targeting framework in 1991, the inflation target itself was revised downward throughout the 1990s and only stabilized at the current range of 2–4 percent in the first quarter of 2001. Finally, the Chilean yield curves are based on a relatively small number of securities, owing to the smaller size of Chilean financial markets,

36. Our data on U.S. macroeconomic data releases remain relatively free of measurement error, however. We consider the response of Chilean interest rates to these U.S. releases, just as we did for Canada in the preceding section.

so that implied forward rates for Chile are generally much noisier than in the United States and Canada, again posing a challenge for empirical analysis.

We report the results of our analysis for Chile in tables 5 and 6. Table 5 reports the response of Chilean interest rates and inflation compensation to domestic economic news. The first column of the table reports the estimated responses of short-term Chilean interest rates to economic news over this period. Only one of our four Chilean macroeconomic data releases—monetary policy announcements—is statistically significant, which is consistent with the idea that measurement error and a shorter sample make estimation difficult. That one statistic is highly significant, however, with a sign and magnitude similar to our estimates for the United States. Moreover, the joint test of the hypothesis that all coefficients in the regression are zero can be rejected at the 1 percent significance level. We thus have evidence that our analysis still has power despite the limitations of the data. Nevertheless, in contrast to the behavior of Chilean short rates, neither far-ahead forward nominal rates nor inflation compensation respond significantly to Chilean monetary policy announcements, which suggests some degree of anchoring. The hypothesis that all of the coefficients in these regressions are zero cannot be rejected at any standard level of significance.

In table 6, we address the response of Chilean interest rates to U.S. macroeconomic and monetary policy announcements. A few U.S. statistics are estimated to have significant effects on Chilean short rates, although some of the coefficients (on U.S. nonfarm payrolls and unemployment) have signs that are perhaps puzzling. The joint hypothesis that all coefficients in the short-rate regression are zero is rejected at the 1 percent level. Again, in contrast to short rates, far-ahead forward nominal rates and inflation compensation in Chile respond to almost no U.S. macroeconomic data releases, with the exception of the U.S. unemployment rate release and perhaps U.S. monetary policy surprises. The hypothesis that all coefficients in the regression are zero is also not rejected at standard significance levels in either case. While the Chilean data are clearly much noisier and more problematic than in the data for more industrialized countries such as Canada, Sweden, the United Kingdom, and the United States, our results for Chile are all consistent with those other countries. The exercise suggests that the commitment of the central bank to a credible long-run inflation objective significantly helps to anchor private sector expectations about long-run inflation outcomes.

Table 5. Chilean Forward Rate Responses to Domestic Macroeconomic News, 2002–05[a]

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in ten years*	*One-year forward real rate ending in ten years*	*One-year forward inflation compensation ending in ten years*
CPI	0.40	1.86	–1.37	3.23
	(0.64)	(0.84)	(–0.53)	(0.97)
Real GDP	0.25	1.10	2.13	–1.03
	(0.30)	(0.38)	(0.62)	(–0.23)
Real GDP, previous quarter	–0.69	1.91	2.83	–0.92
	(–0.49)	(0.39)	(0.49)	(–0.13)
Monetary policy	0.15***	0.22	0.06	0.16
	(3.92)	(1.61)	(0.37)	(0.78)
No. observations	98	98	98	98
R^2	0.16	0.04	0.02	0.02
Joint test p value	0.005***	0.406	0.703	0.773

Source: Authors' computations.

* Statistically significant at the 10 percent level. ** Statistically significant at the 5 percent level. *** Statistically significant at the 1 percent level.

a. The sample is from August 2002 to October 2005, at daily frequency on the dates of macroeconomic and monetary policy announcements. Regressions also include a constant and a year-end dummy that takes on the value of 1 on the first business day of any year (coefficients not reported). Macroeconomic data release surprises are normalized by their standard deviations, so these coefficients represent a basis point per standard deviation response. Monetary policy surprises are in basis points, so these coefficients represent a basis point per basis point response. Inflation compensation is the difference between nominal and real rates. Joint test p value is for the hypothesis that all coefficients (other than the constant and dummy variables) are zero. t-statistics are reported in parentheses.

Table 6. Chilean Forward Rate Responses to U.S. Macroeconomic News, 2002–05[a]

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in 10 yrs*	*One-year forward real rate ending in 10 yrs*	*One-year forward inflation compensation ending in ten years*
U.S. capacity utilization	−0.16 (−0.23)	2.27 (1.02)	−1.06 (−0.39)	3.33 (0.96)
U.S. consumer confidence	−0.05 (−0.08)	−0.59 (−0.27)	−0.02 (−0.01)	−0.57 (−0.17)
U.S. core CPI	0.86 (1.11)	2.12 (0.85)	−4.19 (−1.39)	6.31 (1.63)
U.S. employment cost index	0.78 (0.81)	0.65 (0.21)	4.00 (1.07)	−3.35 (−0.70)
U.S. real GDP (advance)	−0.44 (−0.32)	−4.92 (−1.14)	2.95 (0.57)	−7.87 (−1.17)
U.S. initial jobless claims	−0.36 (−0.93)	0.80 (0.66)	−0.65 (−0.44)	1.46 (0.76)
U.S. NAPM/ISM manufacturing survey	−0.60 (−0.66)	−0.26 (−0.09)	5.23 (1.50)	−5.49 (−1.22)
U.S. new home sales	0.38 (0.80)	0.38 (0.25)	−2.53 (−1.39)	2.92 (1.24)
U.S. nonfarm payrolls	−1.35* (−1.72)	1.55 (0.62)	−3.50 (−1.16)	5.06 (1.30)
U.S. retail sales (excl. cars)	1.68** (2.20)	−2.46 (−1.01)	0.48 (0.16)	−2.94 (−0.77)

Table 6 (continued)

Explanatory variable	*One-year nominal rate*	*One-year forward nominal rate ending in 10 yrs*	*One-year forward real rate ending in 10 yrs*	*One-year forward inflation compensation ending in ten years*
U.S. unemployment rate	3.78***	–8.70***	2.98	–11.68***
	(4.20)	(–3.03)	(0.86)	(–2.61)
U.S. monetary policy	0.25	–0.81	0.67	–1.48*
	(1.46)	(–1.51)	(1.04)	(–1.77)
No. observations	399	399	399	399
R^2	0.10	0.05	0.03	0.05
Joint test *p* value	0.001***	0.234	0.688	0.167

Source: Authors' computations.
* Statistically significant at the 10 percent level.
** Statistically significant at the 5 percent level.
*** Statistically significant at the 1 percent level.
a. The sample is from August 2002 to October 2005, at daily frequency on the dates of macroeconomic and monetary policy announcements. Regressions also include a constant, a year-end dummy that takes on the value of 1 on the first business day of any year (coefficients not reported), and Chilean macroeconomic news releases (coefficients not reported since they are very similar to table 5). Macroeconomic data release surprises are normalized by their standard deviations, so these coefficients represent a basis point per standard deviation response. Monetary policy surprises are in basis points, so these coefficients represent a basis point per basis point response. Inflation compensation is the difference between nominal and real rates. Joint test *p* value is for the hypothesis that all coefficients (other than the constant and dummy variables) are zero. *t*-statistics are reported in parentheses.

4.3 Time-Series Behavior of Forward Rates in Canada, Chile, and the United States

Our analysis in the previous sections focused on the conditional volatility of forward rates in Canada, Chile and the United States, by which we mean the movement of these rates in response to specific data releases. Although we took care to include as many variables as possible and any macroeconomic data release that seemed important, our regressions have nonetheless omitted many factors that influence the daily behavior of interest rates at both the short and long ends of the yield curve. The R^2 values in our regressions are in every case below 20 percent, even for short-term interest rates.[37] Given our argument that the relative responsiveness of forward rates in different countries to macroeconomic data releases and monetary policy announcements is due to different degrees of stability of private sector long-run inflation expectations, one might expect to see that other economically relevant news that we have omitted would lead to a similar contrast in far-ahead forward interest rate behavior across our three countries. In other words, one might expect to see forward rates in the United States that would tend to be more volatile unconditionally as well as conditionally, to the extent that long-run inflation expectations in the United States are unanchored.

Figure 5 presents unconditional time series plots of far-ahead forward nominal rates and inflation compensation for Canada, Chile, and the United States. We find a number of interesting observations. First, far-ahead nominal rates and inflation compensation are not completely stable in any of the three countries. Both high and low frequencies exhibit clear variation, the source of which remains an open question. Possible explanations include the following: high transaction

37. This observation is all the more remarkable in light of the fact that we have restricted our attention in the regression to only those days on which at least one of our right-hand-side variables was nonzero; the R^2 values are even lower (though our coefficient estimates are very similar) if we perform the regression on all days. Thus, even on the days on which important macroeconomic news was released, we can only explain a relatively small fraction of the variance of interest rates at even the short end of the yield curve. One reason for the low R^2 values is that macroeconomic data releases often contain much more information than just the simple headline number that we must focus on in our analysis. For example, monetary policy announcements by the Federal Reserve often contain lengthy statements discussing the motivation for the move and even the future outlook for monetary policy; GDP releases contain information about its various components, which can independently influence private-sector forecasts of future output; and inflation releases contain a detailed breakdown of constituent components, which may independently influence forecasts of future inflation.

costs in Canadian and Chilean markets that drive observed prices away from true shadow values and errors in yield curve estimation resulting from a small number of securities outstanding;[38] time-varying risk or liquidity premiums; variations in financial market perceptions of the central bank's credibility and commitment to its long-run inflation objective; changes in the official inflation target itself (both Canada and Chile lowered their official targets several times in the early 1990s) or perceptions that the central bank's inflation target might change in the future; changes in tax rates or market perceptions that tax rates might change in the future; market perceptions that the central bank's preferred measure of inflation might change in the future; and differences between the consumption deflator of the marginal investor and the price index that is being targeted by the central bank.

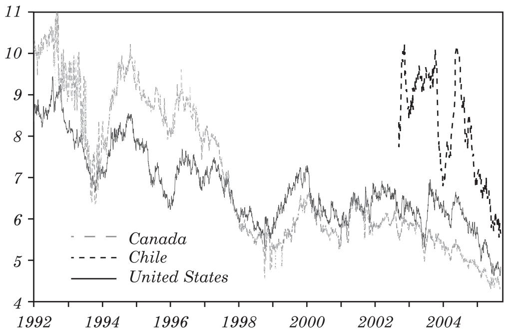
Second, despite the variation in our estimates of far-ahead forward nominal rates and inflation compensation, the Canadian rates have improved spectacularly vis-à-vis the United States. In the first half of the 1990s, far-ahead forward rates in Canada were clearly and consistently higher and more volatile than in the United States. From the late 1990s onward, that situation has completely reversed: far-ahead forward nominal rates and inflation compensation in Canada have been clearly and consistently lower and less volatile than in the United States. This is all the more remarkable considering that liquidity is lower and transaction costs higher in Canada, and the number of outstanding securities with which to estimate a yield curve is much smaller; thus, all else equal, one would tend to expect risk premiums and measurement error to produce more volatile forward rates in Canada. These observations exactly parallel the findings of Gürkaynak, Levin, and Swanson (2006) for the United Kingdom and Sweden. The sample period for our Chilean data is shorter, but it also shows a remarkable fall in these far-ahead forward rates over time, bringing them toward levels that are becoming increasingly comparable to those in the United States.

Third, inflation targeting by itself is not a silver bullet that suddenly lowers and stabilizes far-ahead forward nominal rates

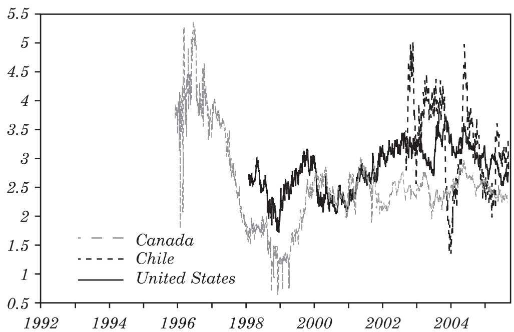
38. As mentioned in the preceding sections, Chile has only a few nominal and indexed government bonds outstanding, and Canada has only a few highly illiquid indexed government securities outstanding. Thus, estimates of forward rates in these two countries can be noisy, particularly in Chile and in the early years of the Canadian indexed market, when there were only two bonds outstanding and their liquidity was very low. (A third Canadian real bond was introduced in 1999 and liquidity in that market has increased steadily over time).

Figure 5. Time Series Plots of Forward Nominal Rates and Inflation Compensation

A. One-year forward nominal rate ending in ten years (percent)



B. One-year forward inflation compensation ending in ten years (percent)



Source: Authors' computations.

and inflation compensation. Canada officially adopted an inflation-targeting framework in February 1991, but the real gains in far-ahead forward rates and inflation compensation seem to have come gradually. Why this is so remains an open question, but it may be due partly to the fact that, although Canada adopted a formal inflation-targeting framework in 1991, the official inflation target was revised lower on several occasions in the early 1990s. One would hardly expect long-term inflation expectations to be anchored around the central bank's target if that target itself were perceived by markets to be in transition to an unspecified long-run level. Thus, the true date of adoption of a fixed long-run inflation target in Canada might be identified as 1995, the date at which the current range of 1–3 percent was adopted and regarded as likely to persist (Mishkin and Schmidt-Hebbel, in this volume, make this point for a number of inflation-targeting adopters).[39] In addition, the initial announcement of an inflation-targeting regime in Canada and the initial announcement of the 1–3 percent target may have been regarded with some skepticism by financial markets, and only gradually did the feasibility of—and the central bank's commitment to—the new targeting regime become clear. These factors may also help explain why far-ahead forward nominal rates and inflation compensation in Chile remain fairly volatile and have exhibited somewhat of a downward trend in the past few years.

Finally, the figure provides direct evidence against the critique by Ball and Sheridan (2003) that there are no visible benefits from inflation targeting once initial conditions and mean reversion are taken into account. Ball and Sheridan's argument would predict that Canada, which began from high levels of inflation expectations in the early 1990s, would tend to converge back toward the levels in the United States over the 1990s. In contrast to this prediction, however, we find that inflation expectations in Canada actually overtake those in the United States in 1997 and then outperform the United States for the next eight years. This is a much stronger performance than can be accounted for simply by a tendency for reversion to the mean.

5. Conclusions

As in Gürkaynak, Sack, and Swanson (2005) and Gürkaynak, Levin, and Swanson (2006), we find that U.S. long-term forward nominal interest rates and inflation compensation are excessively sensitive to macroeconomic data releases and monetary policy announcements. In contrast, we find that long-term nominal interest rates and inflation compensation in Canada display much less sensitivity to economic news, while the unconditional volatility of these series over the past decade has been markedly lower than in the United States. These results are consistent with the findings of Gürkaynak, Levin, and Swanson (2006) for Sweden and the United Kingdom, two countries that have also maintained explicit inflation targets in recent years

39. The adoption of a target range for inflation (as opposed to a point) is not, in itself, a reason for variability of long-term inflation expectations, since the optimal monetary policy is always to aim for the midpoint of the range, as noted previously in this paper and discussed in detail in Orphanides and Wieland (2000).

In the case of Chile, the available sample period is fairly short and only a limited set of macroeconomic news releases are readily available. Nevertheless, our regression analysis does not indicate any excess sensitivity of far-ahead forward interest rates and inflation compensation, which is consistent with the hypothesis that inflation targeting in Chile has been reasonably successful in anchoring long-run inflation expectations. The unconditional volatility of these series, however, appears to be much higher in Chile than in either Canada or the United States, perhaps underscoring the extent to which the Chilean economy is still in the process of converging to the economic and financial conditions of the more industrialized economies. In particular, only a small number of Chilean government securities are actively traded in bond markets, and the yields on these securities may be quite sensitive to variations in liquidity and other market frictions. While not entirely conclusive, these results suggest that the presence of a transparent and credible inflation objective can play an important role in anchoring long-run inflation expectations in both emerging market economies and industrialized countries.

Our findings suggest that the potential welfare gains from reduced bond market volatility would be an important subject for future research. Although we have not demonstrated any such welfare gains in this paper, existing macroeconomic and finance theory identifies several strong possibilities: for example, less persistent deviations of inflation from target in the short and medium run as a result of firmer anchoring of expectations at the long end (Woodford, 2003); a greater ability of the central bank to control inflation in the short and medium run (ibid.); less volatile long-term nominal interest rates and lower risk premiums on nominal rates, which would improve the efficiency of investment decisions (Ingersoll and Ross, 1992); and a reduced chance of either a 1970s-style expectations trap for inflation (Albanesi, Chari, and Christiano, 2003) or an imperfect-information-driven inflation scare (Orphanides and Williams, 2005). To the extent that these benefits are important in practice as well as in principle, adopting a more explicit inflation objective could improve U.S. economic performance and U.S. monetary policy even beyond the successes of the past twenty years.

Appendix

Data on U.S. macroeconomic statistical releases and forecasts were obtained from Money Market Services (MMS) through July 2003, when that company merged with a larger financial institution. Beginning in December 2003, the same survey was produced again by Action Economics (AE). Both data sets can be obtained from Haver Analytics at www.haver.com. From August through November 2003, we fill in the holes in the MMS/AE survey data using the releases and forecasts reported by Bloomberg Financial Services. For additional details about individual macroeconomic series, see Gürkaynak, Sack, and Swanson (2003).

We obtained data on Canadian macroeconomic news releases and financial market expectations of those releases from two sources: Money Market Services and Bloomberg, as discussed in section 4. When those data sets overlap, they agree very closely. Between these two data sources, we have data on Canadian capacity utilization, the consumer price index, core consumer price index, employment, real GDP, retail sales, the unemployment rate, and wholesale trade. Most of these series go back to 1996, and a few go back even farther. To measure the surprise component of Canadian monetary policy announcements, we obtained the dates of changes in the Bank of Canada's target overnight interbank rate back to 1995 from the Bank of Canada's web site, and we measured the surprise component of these changes using the change in the three-month Canadian Treasury Bill on the dates of these monetary policy changes. The exact statistics we use, including Bloomberg and MMS mnemonics for those series, are reported in table A1.

Table A1. Data Sources for Canada

Series	*Data Source*	*Mnemonic*[a]	*Notes*
Capacity utilization	MMS	{L,D,M}156CU	Level, percent
Consumer price index	Bloomberg	cacpiyoy	Year-on-year change, percent
Core CPI	MMS	{L,D,M}156CPXY	Year-on-year change, percent
Employment	MMS	{L,D,M}156ED	Month-on-month change, thousands
Real GDP	MMS	{L,D,M}156GPA	Quarter-on-quarter change, percent
Retail sales	Bloomberg	carsmom	Month-on-month change, percent
Unemployment rate	Bloomberg	caunemp	Level, percent
Wholesale trade	Bloomberg	cawtmom	Month-on-month change, percent
Monetary policy	Bank of Canada		Authors' calculations from policy change dates and three-month Canadian Treasury bill rate

Source: Authors' computations.

a. This column reports the mnemonic used in the Bloomberg database for series obtained from Bloomberg and from the MMS database for series obtained from Money Market Services.

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