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The evolution of inflation expectations in Canada and the US

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Abstract. We model inflation forecasts as monotonically diverging from an estimated long-run anchor point towards actual inflation as the forecast horizon shortens. Fitting the model with forecaster-level data for Canada and the US, we identify three key differences between the two countries. First, the average estimated anchor of US inflation forecasts has tended to decline gradually over time in rolling samples, from 3.4% for 1989–1998 to 2.2% for 2004–2013. By contrast, it has remained close to 2% since the mid-1990 for Canadian forecasts. Second, the variance of estimates of the long-run anchor is considerably lower for the panel of Canadian forecasters than US ones following Canada's adoption of inflation targets. And third, forecasters in Canada look much more alike than those in the US in terms of the weight that they place on the anchor. One explanation for these results is that an explicit inflation-targeting regime (Canada) provides for less uncertainty about future monetary policy actions than a monetary policy regime where there was no explicit numerical inflation target (the US before 2012) to anchor expectations.

Résumé. L'évolution des anticipations d'inflation au Canada et aux États-Unis. On présente un modèle de prévision d'inflation en tant que divergence monotone d'une estimation d'un point d'ancrage à long terme vers le taux d'inflation réalisé à mesure que l'horizon de prévision diminue. Ajustant le modèle aux données canadiennes et américaines au niveau des spécialistes en prévision, on identifie trois différences entre les deux pays. D'abord, la moyenne estimée du point d'ancrage pour l'inflation aux États-Unis a eu une tendance à décroître graduellement dans le temps dans les échantillons successifs – de 3,4 % pour la période 1989–1998 à 2,2 % pour la période 2004–2013. Au Canada, ce taux est demeuré près de 2 % depuis le milieu des années 1990. Ensuite, la variance des estimations du point d'ancrage à long terme est considérablement plus faible pour le panel des experts canadiens que pour ceux des États-Unis depuis l'adoption de cibles d'inflation au Canada. Finalement, les spécialistes en prévision du Canada ressemblent grandement à ceux des États-Unis pour ce qui est de la valence qu'ils donnent au point d'ancrage. Une explication pour ces résultats est que le régime de cible explicite d'inflation (Canada) entraîne une moins grande incertitude à propos des actions futures au plan de la politique monétaire

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que le régime américain où il n'y a pas de cible explicite et numérique pour l'inflation (États-Unis avant 2012) pour ancrer les anticipations.

JEL classification: E31, E58

1. Introduction

Central bank transparency can improve the transmission of monetary policy. A clear understanding of how monetary policy will respond to shocks might reduce uncertainty about future policy responses. One measure that could be helpful is having a credible numerical objective for monetary policy. Combined with a model of the economy, this may enable economic agents to accurately predict how monetary policy will respond to economic developments, thereby contributing to macroeconomic stability.

One way to assess the success of a monetary policy regime in this regard is to examine the behaviour of macroeconomic forecasts. If forecasters seek to produce timely, accurate estimates of future outcomes, then these should reflect beliefs about future monetary policy decisions. We carry out such an assessment here.

We focus on Canada and the US, over the 1989–2013 period, during which their respective monetary policy frameworks have evolved in quite different ways. Canada was one of the earliest adopters of a formal inflation target, although the precise nature of that target has evolved over its life. Further, the Bank of Canada adopted an inflation target at a time when it was seeking to lower the inflation rate. In contrast, the US did not have a specific numerical target for inflation before 2012, and its adoption of a target occurred following a period when inflation had been unusually low.[1]

In comparing inflation forecast behaviour in Canada with the US, this paper is related to a rich literature examining the effects of inflation targeting on inflation forecasts. A straightforward way to do this is to focus on mean forecasts values across a panel of forecasters, especially at longer horizons. For example, Levin et al. (2004) find that longer-term inflation forecasts are less correlated with lagged inflation in five economies with explicit inflation targets than in seven economies without, and Davis (2014) finds that the sensitivity of inflation expectations to oil price and inflation shocks has declined in inflation targeting economies. Such forecasts are available at horizons of up to 10 years. Figure 1 contains 6 to 10 year ahead mean forecasts for Canada and the US from Consensus Economics. The date on the horizontal axis is when the forecasts were made.

Even from a cursory glance, it is clear that long-run inflation expectations in the two countries have behaved quite differently from each other. Those in the US have trended down over most of the sample period but have remained relatively

1 See section 3 of the working paper version of this paper (Yetman 2015) for a detailed discussion of the evolution of the monetary policy frameworks in Canada and the US.

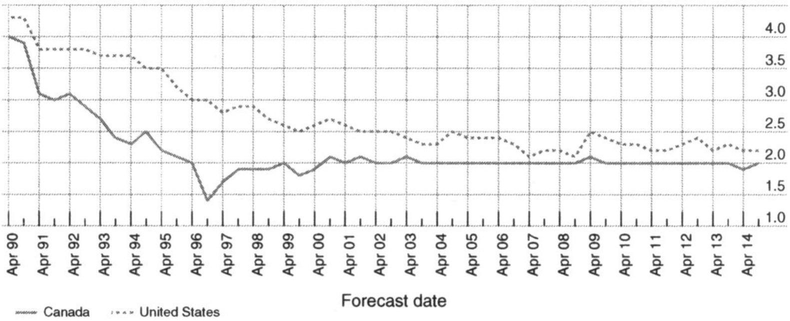

FIGURE 1 Forecasts of long-run inflation, 6–10 years ahead
SOURCE: Consensus Economics©

volatile throughout. By contrast, those for Canada fell rapidly early in the sample period, after which they have remained close to 2%.

However, looking only at these average forecasts ignores a lot of relevant information. First, they are available only twice a year (whereas shorter term forecasts are available at a monthly frequency). Second, and more important, only average forecasts are available. These suppress useful information on the degree of dispersion across the panel of individual forecasters. Using shorter-term forecasts, for example, Johnson (2002) compares inflation forecasts for five inflation targeters with six non-targeters and finds that inflation targeting lowered the level of inflation expectations but had no effect on the variability of expected inflation. Cecchetti and Hakkio (2009) argue that the effect of inflation targeting on the dispersion of inflation forecasts has been small. And Capistrán and Ramos-Francia (2010) find that longer-term expectations are less dispersed in inflation targeting emerging market economies, with a cumulative effect that builds over the first three years following the adoption of inflation targets, but find little effect for advanced economies.[2,3]

However, there are limitations with how much of the existing literature uses the multiple horizons for which these shorter-horizons forecasts are available. One approach is to cluster different forecast horizons into two groups: for example "current year" and "following year." Variation across horizons may then be treated as a form of seasonality, with horizon dummy variables included in the estimation. Another, in Siklos (2013), uses a weighted average of two fixed-

2 Capistrán and Ramos-Francia (2010) define inflation expectations as "anchored when individual expectations with a forecast horizon equal to or greater than the central bank's control lag are at or very close to the inflation target, even if inflation [...] is not at or close to the target." The model that we estimate can be interpreted as a generalization of this idea in the sense that the control that the central bank has over inflation outcomes is a monotonically increasing function of the forecast horizon, rather than assuming complete control over inflation beyond some specific horizon.

3 Note, also, Autrup and Grothe (2014), Galati et al. (2011), Nautz and Strohsal (2015) and Strohsal and Winkelmann (2015), who assess inflation expectations anchoring based on the behaviour of break-even inflation rates.

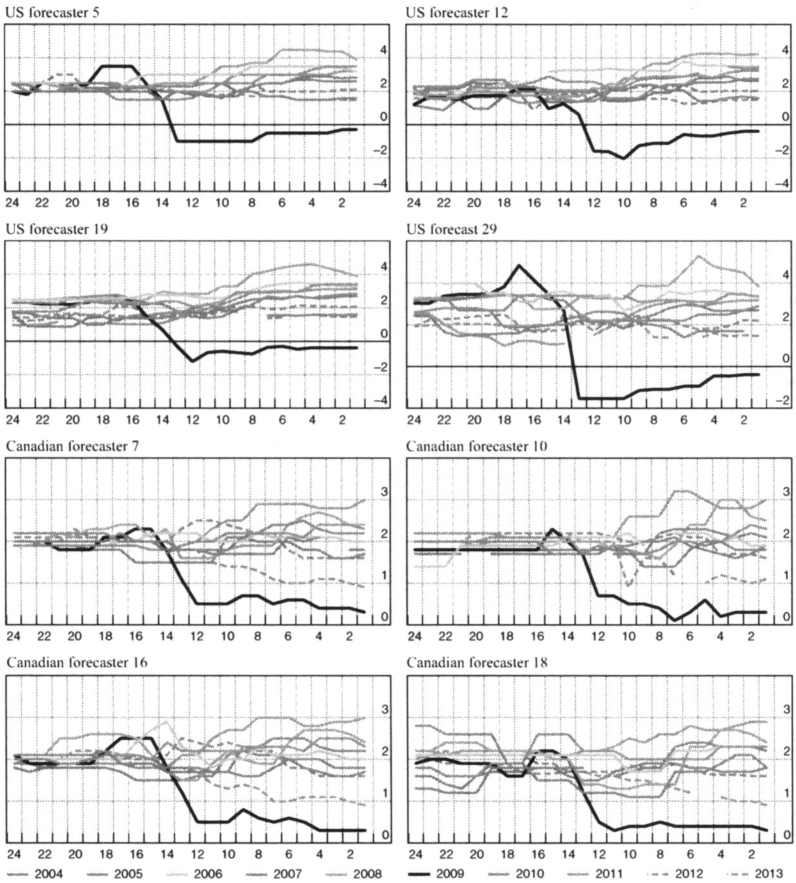

FIGURE 2 Forecasts of headline inflation at different horizons
NOTE: Horizontal axis corresponds to the number of months before the completion of the year being forecast, i.e., 24 indicates forecasts made in January for the following calendar year.
SOURCE: Consensus Economics©

event forecasts to approximate fixed-horizon one-year ahead forecasts.[4] These approaches may not capture the dynamics seen across forecast horizons in the forecast data.

Figure 2 displays the forecasts from the four most frequent forecasters for the 2004–2013 period from each country in our sample. The horizontal axes are the forecast horizons, which run from $h = 24$ (forecasts made at the start of January, 24 months before the completion of the calendar year being forecast, the longest horizon at which Consensus Economics makes individual forecaster-level forecast data available) to $h = 1$ (made in December of the calendar year being forecast). The forecasts across the different years for a given forecaster are much more

4 See Mehrotra and Yetman (2014) for a more detailed critique of this approach.

similar to each other at longer horizons than at shorter ones. They increasingly diverge as the horizon shortens, and at short horizons they look a lot like the distribution of inflation outcomes (not shown).

We focus on the forecasts of individual forecasters at horizons of up to two years and examine how these vary over time using a framework that can capture these dynamics. We apply the empirical model first introduced in Mehrotra and Yetman (2014) to these data. Inflation forecasts for each forecaster are modelled as being the weighted sum of two components: a long-run forecaster-specific anchor that is estimated and the latest available actual inflation rate at the time that the forecast was made, with the weights summing up to one. The weight on the anchor is modelled with a flexible decay function, so that inflation forecasts monotonically diverge from their estimated anchor towards actual inflation as the forecast horizon shortens.

When we compare the distributions of estimates of both the long-run anchors and the decay paths between Canada and the US, we identify three important differences between the two countries:

(i) In rolling samples, the average estimated anchor declines gradually over the sample period for US forecasters. By contrast, it declined early in the sample period for Canadian forecasters and then remained at its new low level, very close to the midpoint of the announced inflation target of 2%.
(ii) The variance of estimates of the long-run anchor for US forecasters has remained at roughly the same level over time. For Canadian forecasters, it declined from levels comparable to those for the US early in our sample period to around one quarter of the US level by the end.
(iii) And the distribution of estimated weights on the long run anchor varies between the countries, especially later in the sample. Whereas the decay functions look similar for Canadian forecasters, they are bimodal for US forecasters. For the latter country, some forecasters place a high weight on the inflation anchor at all horizons beyond 12 months, while for others the weight decays more uniformly.

In the next section, we outline the methodology and describe the data. Section 3 reports the results, and section 4 concludes.

2. Methodology and data: Forecaster-level data

2.1 Functional form

We adopt the same parsimonious framework for fitting inflation forecasts used in Mehrotra and Yetman (2014) that fully utilizes the multiple-horizon dimension of the available forecast data. The forecast of inflation for year t made at horizon h, denoted $f(t, t-h)$, is assumed to follow:

$$f(t,t-h)=\alpha(h)\pi^{*}+[1-\alpha(h)]\pi(t-h)+\varepsilon(t,t-h). \tag{1}$$

In (1), h is measured in months before the end of the year that is being forecast and π^* is the level that long-run inflation expectations are anchored to, which will be estimated. $\pi(t-h)$ is the level of actual inflation observed at the time when the forecast is made and $\varepsilon(t,t-h)$ is a residual term. To correct for the publication lag in inflation data, we use the 12-month growth rate in monthly CPI lagged by one month as the actual inflation rate. $\alpha(h)$ denotes a decay function, which we model as:

$$\alpha(h)=1-\exp\left(-\left(\frac{h}{b}\right)^{c}\right). \tag{2}$$

This has the desirable features that inflation expectations are equal to the anchor at sufficiently long horizons ($\alpha(\infty)=1$), moves increasingly towards inflation outcomes as the horizon shortens ($\alpha'(h)<$), converges to actual outcomes when the horizon is zero ($\alpha(0)=0$) and provides for a wide variety of possible decay paths for different values of b and c.

The variance of the residual in (1) is modelled using a flexible functional form that allows it to vary across the forecasting horizon with minimal restrictions:

$$V(\varepsilon(h,t))=\exp(\delta_0+\delta_1 h+\delta_2 h^2). \tag{3}$$

We also allow for forecasts for the same inflation outcome made at two different horizons, h and k, to be more highly correlated the closer the horizons are, by assuming that:

$$Corr(\varepsilon(t,t-h),\varepsilon(t,t-k))=1-\phi_1|h-k|-\phi_2(h-k)^2. \tag{4}$$

2.2 The data

Mehrotra and Yetman (2014) applied the above framework to median inflation forecasts across 44 economies. The innovation in this paper is that we apply the framework at the forecaster level. We estimate the model outlined above for Canada and the US, forecaster-by-forecaster and examine the distribution of the estimates across the different forecasters. As we will show, not all forecasters are alike, but this variation is likely to be masked when summary statistics such as medians are examined, instead of looking at the underlying forecaster-level data.

In applying our framework to inflation forecast data from Consensus Economics, we first need to identify forecasters.[5] Ideally, we'd like to identify the individuals making the forecasts and, if they were to move from one organisation to another, follow them in our sample. However, this is impractical since there is no available database that identifies the key individual(s) behind the forecasts. Instead, we follow institutions over time, under the assumption that forecasts from a given institution are likely to be made by similar people, using similar methods, from one month to the next. We also merge forecasters with different names where we are able to find evidence to find evidence to support this: see the appendix for a detailed explanation.

5 Dovern et al. (2015) follow a similar process to match forecasters in their study of GDP forecasts.

Consensus Economics starts collecting forecasts for calendar-year inflation outcomes in the CPI in January of the preceding year ($h = 24$). They collect these forecasts each month until December of the year being forecast ($h = 1$), for a total of 24 monthly forecasts of the same outcome.

3. Results

For many of the forecasters, the panel of forecasts is unbalanced, as is clear from tables A1 and A2, and is visible as gaps in the series plotted in figure 2. We take explicit account of this in our estimation by setting the contribution to the likelihood function to zero for missing observations.

The model is estimated by maximum likelihood, forecaster-by-forecaster, using 10-year rolling samples (where years are defined in terms of the inflation rate being forecast, rather than when the forecasts are made) to allow for the degree of anchoring to evolve over time.[6] We include all rolling samples where more than 50% of the possible 240 observations are present in the sample for a forecaster and there is a forecast available for at least one horizon of both the first and last years of the rolling sample. We consider 40 different possible starting values for each sample and maximize the likelihood function using the hill-climbing method of Broyden, Fletcher, Goldfarb and Shanno (see Shanno 1985 for details) for each, until the estimates converge. We then choose the estimates with the highest loglikelihood function value where the parameters of the decay function and the inflation target are identified ($b > 0, c > 0, V(b) > 0, V(c) > 0, V(\pi^*) > 0$). In most cases, a majority of the starting values considered lead to virtually identical parameter estimates; in all cases, our selection criteria lead to unique estimates for each rolling sample—forecaster combination.

3.1 The estimated inflation anchor

We first present results on the estimates of the inflation anchors across our panels. For each rolling sample, we have one estimate of the inflation anchor for each forecaster.[7]

Figure 3 provides the mean, together with 95% confidence bands (based on the standard deviation of estimates for each forecaster), of the inflation anchor estimates for the US. There is a minimum of 11 forecasters in each rolling sample, with the number of forecasters increasing over early samples and stabilising at 22 to 24 forecasters beginning 1993–2002.

The results indicate that the average estimated inflation anchor has come down steadily over time, from around 3.5% to just above 2% by the final rolling sample. The width of the confidence bands around the forecasts, a measure of uncertainty

6 The anchor (along with other parameters) is implicitly assumed to be constant within the 10-year rolling sample. Provided this is slow moving, as appears to be the case, rolling samples may be adequate to capture variation over time.

7 Detailed estimation results are available from the CJE online archive at `economics.ca/cje/en/archive.php`.

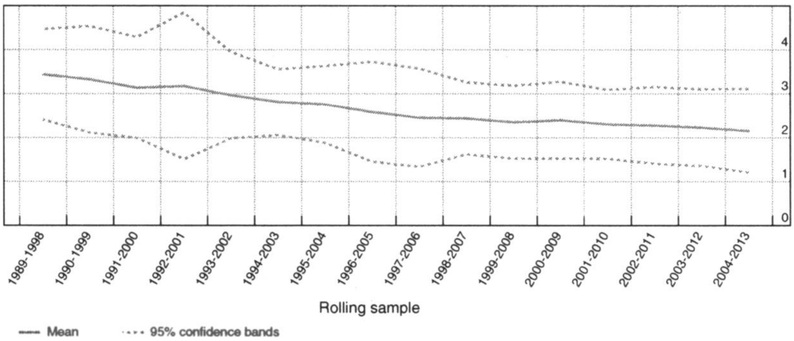

FIGURE 3 Estimated US inflation anchors: Mean and 95% confidence bands
SOURCE: Author's calculations

about the future inflation rate, has remained approximately constant over the sample.

We demonstrate the robustness of the results in a number of different ways. We consider dropping the forecasters with the highest and lowest anchor estimates from each rolling sample. We also drop all forecasters who are present for only a small number of rolling samples, to reduce the influence of forecasters who are not always present in the panel, with two different thresholds. We first exclude all forecasters who are present for less than four rolling samples, followed by 10 rolling samples. In the latter case, our sample consists of 8 to 17 forecasters, six of whom are present in every sample. We also consider the median as an alternative summary measure of the level of the inflation anchor and the inter-quartile range in place of the standard deviation.

A full set of results for all of these measures is given in figure 4. In all cases, the results are consistent with those reported above: the inflation anchor has trended down, while uncertainty about the inflation anchor has no clear trend over time, with the exception that there has been an increase in the degree of uncertainty across all of these measures in the final few rolling samples.

Turning to Canada, figure 5 is the analogue of figure 3. For the Canadian sample, there are 10 to 12 forecasters present in the first 10 rolling samples, increasing to 16 by the final sample. As with the US, the inflation anchor has declined by close to one percentage point over the sample period. But, in contrast to the US case, nearly the entire decline occurred early in the sample. Indeed, by the 1993–2002 rolling sample, the mean of the anchor estimates was within 0.2 percentage points of the midpoint of the Bank of Canada's inflation target, where it remains for all later estimates.

There is also a substantive difference between the variance of the inflation anchor estimates for Canada relative to the US. Whereas in the case of the US the standard deviation remained a similar magnitude across all rolling samples, for Canada there is a significant decline, concentrated in the early part of the sample. For the earliest rolling samples, the level of the standard deviation is

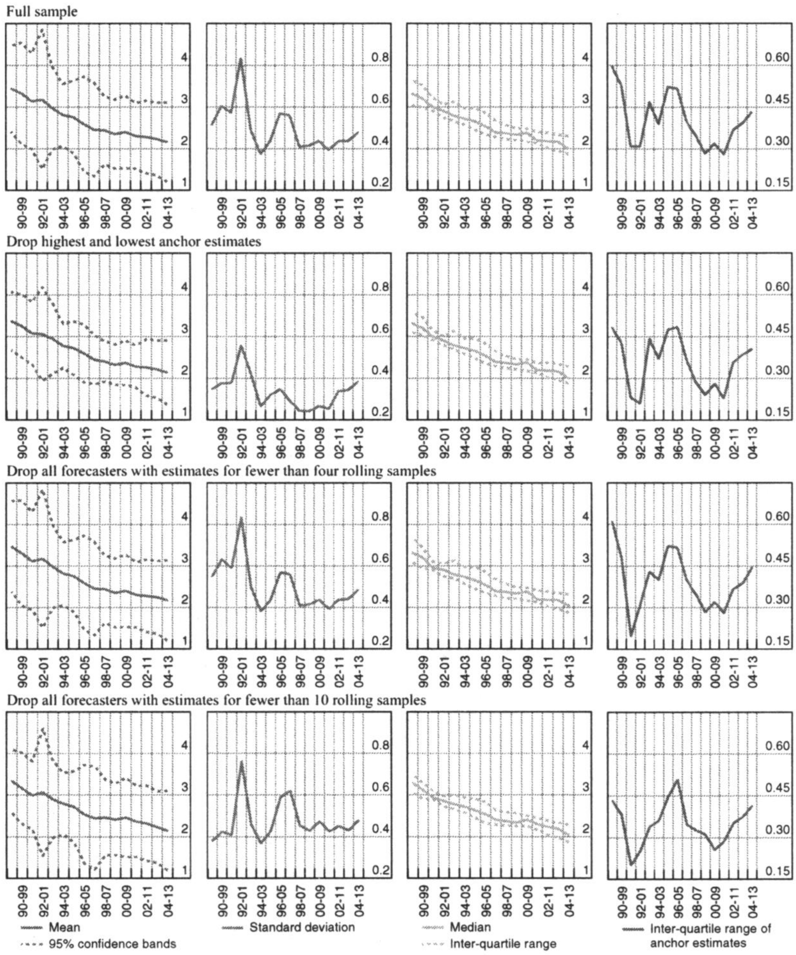

FIGURE 4 Estimated US inflation anchors: Robustness checks
NOTES: From the left: The first column contains the average estimated inflation anchor and 95% confidence band across all forecasters for each rolling sample. The second column contains the standard deviation of estimated inflation anchors. The third column contains median estimated inflation anchors and inter-quartile ranges. The final column contains inter-quartile ranges.
SOURCE: Author's calculations

similar for both countries. But in Canadian samples, it falls rapidly, by a factor of around five, over the first six rolling samples and continues to decline after that, until it is around one quarter of the level for the US by the final rolling sample.

Figure 6 displays robustness checks for Canada. Again, the main results are not sensitive to the different trimming criteria we consider for selecting forecasters or alternative summary measures of the level or variability of anchor estimates.

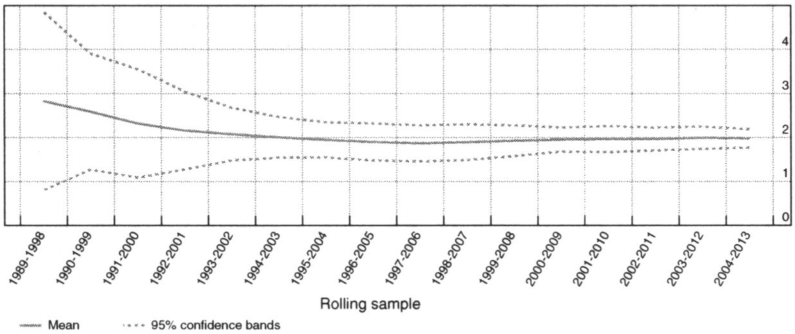

FIGURE 5 Estimated Canadian inflation anchors: Mean and 95% confidence bands
SOURCE: Author's calculations

The only exception to this story is that a continuing decline in the measure of anchor uncertainty in the latter half of the rolling samples is not a robust feature of the results.

Another difference between the two countries is apparent if we focus on the very end of the sample. The final rolling sample contains many observations from the recent financial crisis period, when policy rates have been close to zero. In the Canadian case, the confidence band and inter-quartile range continued to narrow or remained stable. In contrast, all measures of the variability of the estimated anchors for the US widened.

One limitation with the comparison of results so far is its qualitative nature, especially in light of the fact that it is based on estimated coefficients which are subject to random variation. We therefore further examine the difference between the estimates across the two economies statistically. First we take the estimates of π^*, for each forecaster and rolling sample, and regress this on fixed effects for each rolling sample and a dummy that is equal to one for US forecasters and zero for Canadian forecasters. We run a Prais–Winsten regression, allowing for both cross-sectional and serial correlation in the errors.[8]

The estimated coefficient on the dummy is 0.57, indicating that estimated inflation anchors for US forecasters are over half a percentage point higher than those for Canadian forecasters in our sample. This economically large difference is also highly statistically significant, with a p-value of 0.00.[9]

One possibility is that our results reflect, at least in part, reverse causality: changing inflation dynamics are driving changes in the behaviour of inflation forecasts. Clark and Nakata (2008), Stock and Watson (2007), Chowdhury et al.

8 We use the "xtpcse" function in Stata, with the "pairwise" option, and allow the AR(1) coefficient to vary by forecaster. The panel is incomplete and sufficiently sparse that the variance–covariance matrix cannot generally be identified for the full panel. Reported results are therefore based on forecasters who are present in the panel at least 10 times, but are very similar if we use a lower cut-off.

9 All second-stage regression results discussed in the paper are available from the CJE online archive at `economics.ca/cje/en/archive.php`.

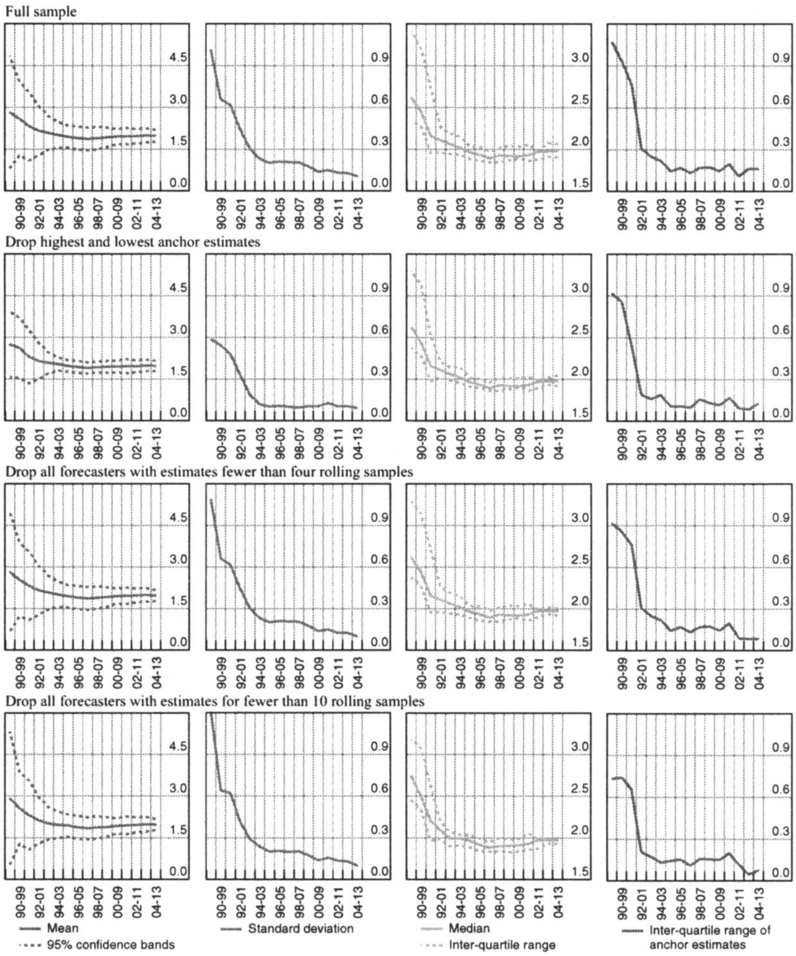

FIGURE 6 Estimated Canadian inflation anchors: Robustness checks
NOTES: From the left: The first column contains the average estimated inflation anchor and 95% confidence band across all forecasters for each rolling sample. The second column contains the standard deviation of estimated inflation anchors. The third column contains median estimated inflation anchors and inter-quartile ranges. The final column contains inter-quartile ranges.
SOURCE: Author's calculations

(2006), Morley et al. (2015) and Bataa et al. (2014) all document changes in inflation dynamics over time. Thus, as a robustness check, we also consider including three variables associated with inflation outcomes during the rolling sample in the regressions: the mean and standard deviation of the inflation rate and inflation persistence, measured as an estimated AR(1) coefficient. None of the three variables is statistically significant, while the coefficient on the dummy remains of similar magnitude (0.51) and statistically significant (p-value 0.02).

In addition, as a final check, given the variation over the sample, we consider simply including year dummies on their own and also interacted with the US dummy. In this case, we find that the estimated anchor is statistically significantly lower for Canadian forecasters than for US forecasters at the 5% level for all rolling samples after the first one (i.e., from 1990–1999 to 2004–2013), with an estimated difference of at least 0.4 percentage points in each sample between 1991–2000 and 2001–2010.[10] Overall, these results reinforce the visual impression from the figures that estimated anchors are lower for Canada than for the US over much of the sample.

3.2 The decay path

Thus far we have focused on estimates of the inflation anchor and demonstrated some important differences in these between Canadian and US forecasters. We also have estimates of how forecasts move away from the anchor, as its importance in explaining forecasts decays with a shortening forecast horizon, to which we now turn.

Our estimates of b and c provide for a wide range of possible decay paths. We use the estimates of these for each forecaster to compute the full decay function path for all horizons and then calculate the median and inter-quartile range of the weight on the inflation anchor, $\alpha(h)$, for different horizons, h. These are displayed for the US in figure 7 and for Canada in figure 8.

For the US, at all except the very shortest horizons, the weight on the anchor has increased over time. At the longest horizons illustrated here, for $h=20$ and $h=24$, $\alpha(h)$ is close to one for the latest rolling sample, versus around 0.6 earlier.

For Canada, it is a similar story, but in general with much less dispersion (as measured by the inter-quartile range). The weight on the anchor increased considerably early in the sample for all $h>10$, consistent with a reduction in the degree of uncertainty about how monetary policy would play out over time.

As with the estimated anchor point, there is also a difference between the two countries that becomes apparent towards the end of the sample, once forecasts made during the recent financial crisis are included in the sample. For US forecasters, the median weight on the long-run anchor declines noticeably at all longer horizons, while there is no obvious change in the case of Canadian ones.

We also examined the difference in the estimated $\alpha(h)$'s empirically for $h \in \{24, 18, 12, 6\}$, paralleling the panel estimation outlined in the previous section. At the 24-month horizon, the estimated weight on the anchor is lower for US forecasters by around 0.07, a difference that is significant (p-value 0.03). At shorter horizons, the estimated difference is larger −0.13, −0.19 and −0.19, respectively) and statistical significant correspondingly rises.

10 Given that there are different numbers of forecasters for different rolling samples and each country, we also considered three different randomly selected samples of 10 forecasters for all rolling samples/countries as a robustness check. With the exception for the coefficient in the second test above becoming insignificant in one of the three cases, results tended to more strongly indicate lower inflation anchors for Canadian forecasters. Later reported results are all robust to this check.

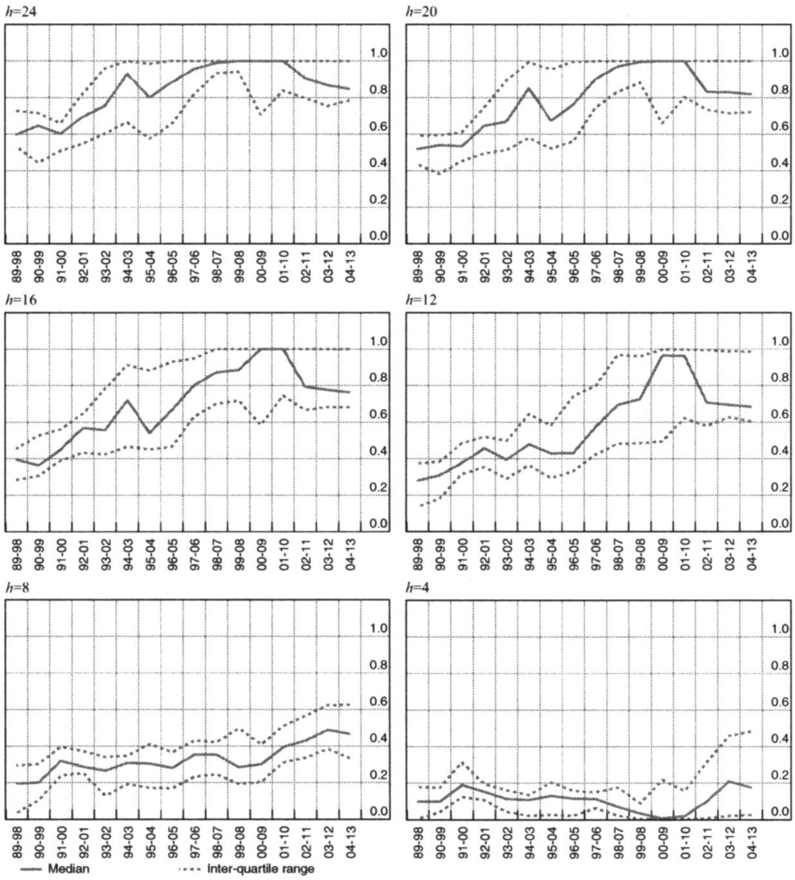

FIGURE 7 Estimated weight on inflation anchor for US forecasters, $\alpha(h)$
SOURCE: Author's calculations

The difference in the $\alpha(h)$'s across countries could in part reflect differences in the underlying dynamics of inflation. But when the mean, standard deviation and persistence of inflation outcomes are all added to the regressions, the coefficient increases in magnitude (to -0.12, -0.20, -0.25 and -0.37 across the different horizons examined), although its statistical significance does deteriorate (p-values are 0.19, 0.06, 0.13 and 0.01, respectively). So at least part of the increase in anchoring may be explained by changes in the dynamics of inflation.

Finally, we examine the decay paths for each of the individual forecasters, since summary measures discussed above may mask some important aspects of forecaster heterogeneity. We focus on three different rolling samples, spread over our sample period. The results for the US are given in figure 9. Of the three samples examined, the decay paths are qualitatively the most similar across the forecasters in the earliest rolling sample, displayed in the top panel. At the longest

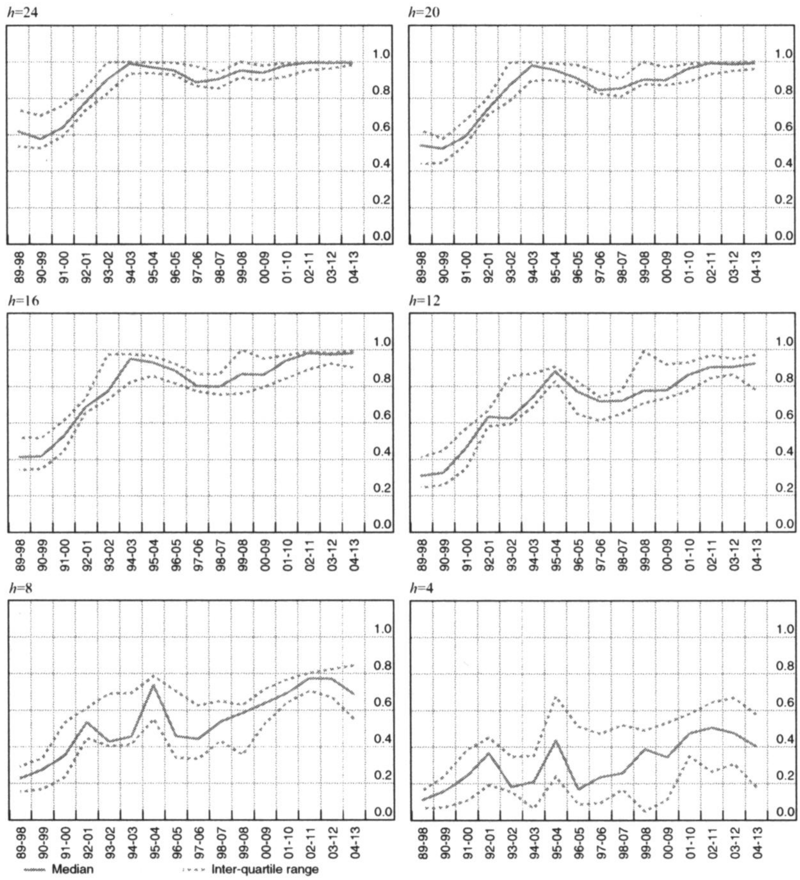

FIGURE 8 Estimated weight on inflation anchor for Canadian forecasters, $\alpha(h)$
SOURCE: Author's calculations

horizons the weight on the anchor varies between 0.5 and 0.9. It then falls as the horizon shortens and lies at around half its initial value when $h = 10$.

Rolling samples in the middle panel, based on estimates over 1996–2005, display much greater dispersion. At the longest horizons, the weight on the anchor varies between 0.4 and 1.0, and the weight on the anchor remains close to 1.0 for the first 12 horizons for a small subset of forecasters. Focusing on the forecasters who are present in both of these earlier panels, the two car manufacturers, Ford and GM (forecasters 10 and 11, respectively), display the greatest increase in the estimated weight on the anchor. At $h = 12$ in the first panel, their weights on the anchor were around 0.5. At the same horizon for the middle panel they both exceed 0.85. For other forecasters present in both panels, the change in the weight on the anchor is relatively smaller. Others for whom, like the car manufacturers, the estimated weight on the anchor is close to one at longer horizons, include a

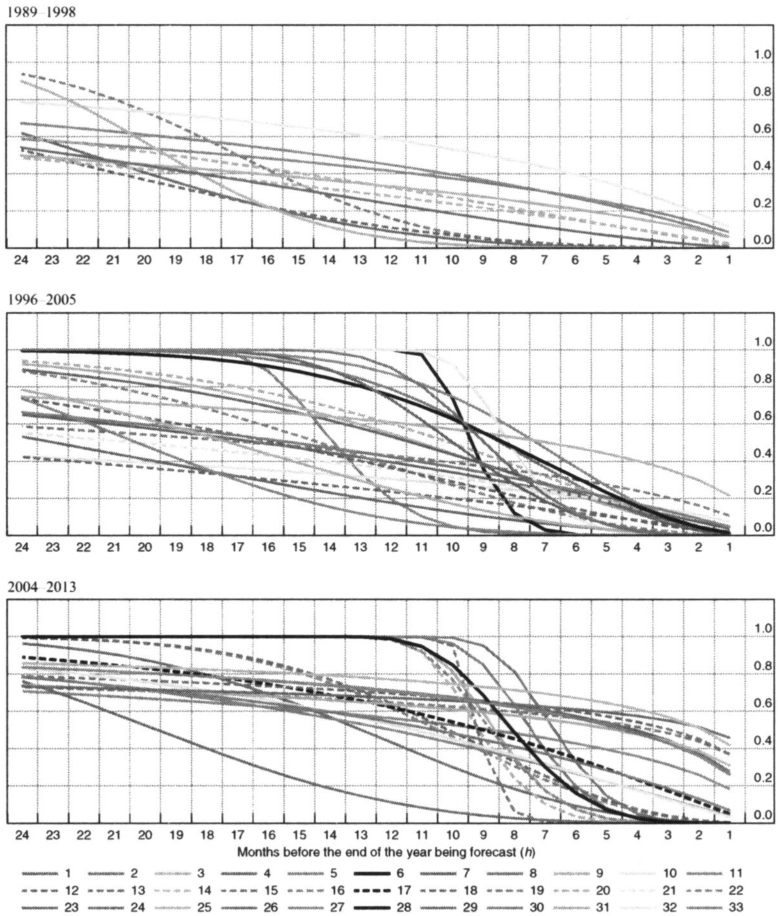

FIGURE 9 US estimated decay paths by forecaster
NOTE: Forecaster numbers correspond to table A1.
SOURCE: Author's calculations

diverse cast of characters: banks (Bank America Corp and its predecessor, Nations Bank, (4); Wells Fargo / Wells Capital (33)), the University of Michigan (30), Standard and Poor's (28) and power management company Eaton Corporation (6).

The lowest panel displays the estimated weights on the anchor for the final rolling sample of 2004–2013. The estimated weights on the anchor at the longest horizons now exceed 0.7 for all forecasters. Except for the conference board (29), an outlier for whom the estimated weight on the anchor falls away quickly, forecasters now fall more clearly into two categories: those whose estimated weight on the anchor remains high for all $h > 12$ and the remainder where the weight on the anchor lies between 0.45 and 0.8 for $h = 12$. There is no apparent

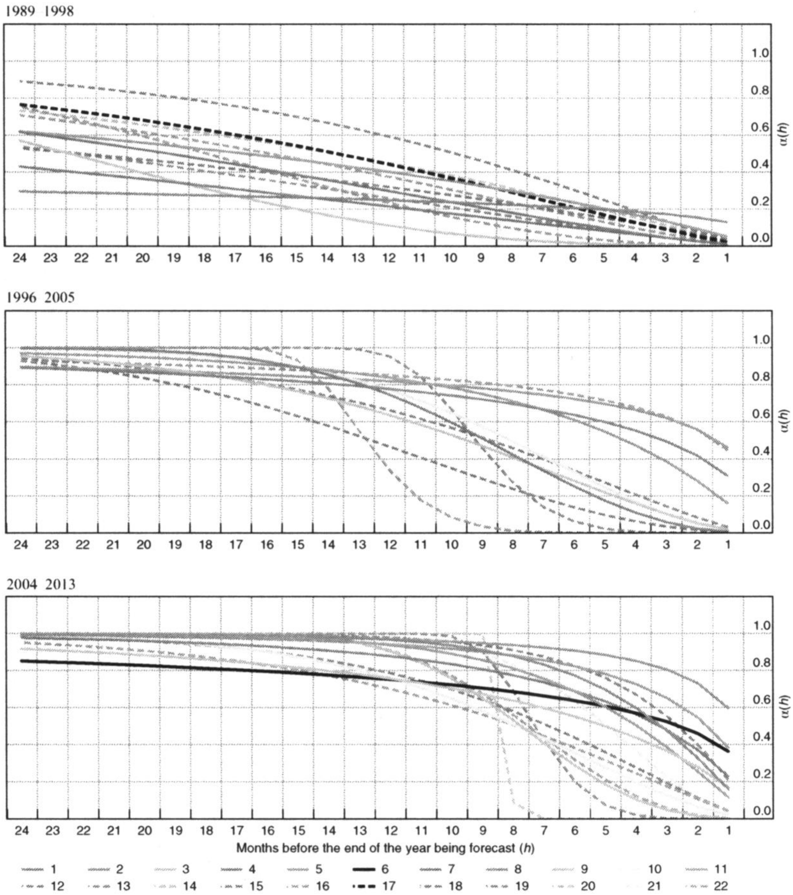

FIGURE 10 Canadian estimated decay paths by forecaster
NOTE: Forecaster numbers correspond to table A2.
SOURCE: Author's calculations

pattern to explain who belongs to each group. The former includes some of those with a high weight on the target in the previous, earlier panel who remain in the sample (6, 30) but does not include others (10, 11, 28, 33), some who were present in the previous sample but with a much lower weight on the anchor for $h > 12$ (Morgan Stanley (20) and IHS Global Insight / Global Insight (31)) and three new-comers (Goldman Sachs (13), Lehman Brothers / Barclays Capital (16) and Oxford Economics (24)). Taken together, the evidence suggests that there is considerable variation in the estimated weight on the anchor across the panel of forecasters across the sample for the US, even at the end of our sample period, that defies simple characterisation.

By way of contrast, figure 10 displays comparable output for Canada. Results from the earliest rolling sample look a lot like those for the same period for the US:

the weight on the anchor varies widely at long horizons and declines smoothly as the horizon shortens. But the later rolling samples, in the middle and lower panels, display much less dispersion, both across time and forecasters, than in the equivalent panels for US forecasters. The estimated weight on the anchor exceeds 0.8 at the longest horizons and it declines in tandem, for all forecasters, as the horizon shortens.

We also examine whether these apparent differences are statistically significant in the following way. We calculate the absolute difference between the weight on the anchor and the median weight across forecasters at horizons $h \in \{24, 18, 12, 6\}$ for each forecaster and then include this as the dependant variable in a panel regression, following the same estimation approach outlined earlier. With time fixed effects (for each rolling sample), the weights on the anchor are further from the median weight for US forecasters by an estimated 0.07 at the longer horizons ($h \in \{24, 18\}$), a result that is highly statistically significant. The estimate is similar if the mean, standard deviation and persistence of inflation outcomes are added to the estimated equation, although it is no longer statistically significant. For $h = 12$, the estimated difference is smaller (0.04; p-value $= 0.01$) but becomes much larger (0.25) and remains statistically significant (p-value $= 0.00$) with the addition of variables reflecting inflation dynamics. Finally, at the shortest horizon we examine ($h = 6$), the estimated difference is negative and marginally significant without the addition of the inflation dynamics variables versus positive and insignificant with them (–0.04 with a p-value of 0.06 vs. 0.10 with a p-value of 0.26). These results indicate that there is generally greater inflation forecast dispersion between US inflation forecasters than between Canadian forecasters, and this can be partly explained by differences in the inflation dynamics between the two countries.

4. Conclusions

In this paper, we have modelled the behaviour of inflation forecasts using a decay function. Inflation forecasts are assumed to monotonically diverge from an estimated anchor towards actual inflation as the forecast horizon shortens. Fitting the data on forecaster-level data for Canada and the US, we have identified three key differences in the behaviour of forecasters between the two countries:

(i) In rolling samples, the average estimated inflation anchor for US forecasters declined gradually over the sample period, whereas for Canadian forecasters, it declined early in the sample and then remained at its new low level, close to the midpoint of the inflation target of 2%.
(ii) The variance across forecasters of the estimates of the inflation anchor has remained at similar levels for US forecasters over time, whereas for Canadian forecasters, it has declined from levels comparable to the US early in the sample to much lower levels.
(iii) The estimates of the weights that the forecasters place on their estimated anchors vary between the countries, especially later in the sample. By the

end of the sample, the distribution for the US looks bimodal, with some forecasters placing a high weight on the inflation anchor at horizons exceeding one year, while the weight for others decays more uniformly and lies between 0.45 and 0.8 at a 12-month horizon. In contrast, there is much less dispersion across Canadian forecasters: the estimated weight on the anchor at a two-year horizon exceeds 0.85 for all forecasters and exceeds 0.7 at a one-year horizon.

What explains these diverse results between the two economies? One possibility is the composition of forecasters. However, while there are some differences between them (there are no car manufacturers among the Canadian forecasters and two among the US, for example), the results do not appear to be driven by this. Even if we were to focus only on banks and financial services firms in the two economies, for example, the systematic differences between them remain.[11] Instead, it seems more likely that differences in the monetary policy regimes are responsible.

Canada adopted an inflation target early in our sample period, with a symmetric target around 2%. Over time forecasters appear to have adjusted their forecasts so that they are strongly anchored at the target level, with only small deviations at forecast horizons exceeding one year. In contrast, the US has not had a clear quantitative target for inflation for most of the sample period, leaving forecasters with no clear anchor for their expectations.

Another important difference between the two economies towards the end of the sample is the degree to which they were affected by the recent crisis, and the extent to which the monetary policy response was constrained by the zero lower bound. For the US, the Federal Funds Rate has been close to 0% since late 2008. For Canada, with the exception of the March 2009 to June 2010 period, the bank rate never dropped below 1%. The zero lower bound may be especially likely to have an impact on inflation expectations if the central bank is thought to lack adequate tools to stabilize inflation, as Ehrmann (2015) discusses. And the use of alternative tools, like quantitative easing, may increase forecaster uncertainty about future inflation rates, as discussed in Hofmann and Zhu (2013). Perhaps not surprisingly, then, there has been a small but noticeable increase in the variability of estimated inflation anchors across US forecasters, and a decrease in the median weight on the anchor at longer horizons, once forecasts made during the crisis are included, changes that are absent from the Canadian estimation results.

Some caveats are in order. First, the results could be subject to reverse causality. Changes in inflation outcomes could feed changes in the behaviour of inflation expectations. However, our second-stage regressions, where we include the mean, standard deviation and persistence of inflation outcomes for the relevant period as explanatory variables, should control for this. Our results are generally robust to

11 At the 12-month horizon for 2004–2013, for example, for four out of nine US banks and financial services firms, the estimated weight on the target exceeds 0.7. The comparable ratio for Canada is 10 out of 10.

their inclusion. Second, there are differences in the monetary policy frameworks between the two countries that could cloud any comparison between them. For example, the inflation target for the US is specified in terms of the PCE price index, whereas that for Canada is in terms of the CPI index. In both countries, we have used the CPI for our analysis, since that is the variable for which we have forecasts. However, there has been a systematic gap between the PCE deflator and the CPI price index in the past of 0.7 percentage points in the 1990s (falling to 0.3 percentage points for 2004–2013), which should be taken into account when interpreting the estimated anchor relative to any published target. Further, to the extent that PCE and CPI dynamics differ, it would be possible for inflation expectations to be anchored more strongly in terms of one variable than the other.

Looking forward, to the extent that the publication of an explicit long-run goal for inflation by the US central bank acts like an inflation target, our results suggest that inflation forecasts for the US might be expected to become more strongly anchored to 2% and the variation seen across forecasts to decline, at least once policy rates are no longer constrained by the zero bound. Recently, Detmeister et al. (2015) reported some results consistent with this. In addition, recent consensus forecasts (made in July 2015) provide us with evidence consistent with these predictions. At the current juncture, US inflation rates are well below the stated target. Average inflation expectations for 2015 of 0.2% reflect this. However, those at a longer horizon indicate a return to near target levels next year (the average forecast for 2016 is 2.2%). This is consistent with our model of forecasts being more strongly anchored the longer the forecast horizon. Additionally, the standard deviation of those longer horizon US forecasts (0.26%) almost matches that for Canada (0.29%), in contrast to past experience.

Appendix: Forecaster identification

One key challenge is identifying forecasters. Taking the full panel of US forecasts of inflation for the 1989–2013 sample period, there are 133 distinct institutions listed. However, this is likely to overstate the number of distinct forecasting identities, since name changes due to reorganisations, mergers and acquisitions may result in little change to the underlying forecasting process. Thus we combine different names wherever two conditions are satisfied: (i) there is plausible evidence that the underlying entities are the same, via corporate websites, news stories or elsewhere and (ii) the timing of the name change lines up with the departure and arrival of the associated names from the panel. For example, Amoco Corporation was present in the panel from the beginning of the sample until February 1999 and BP Amoco starting in March 1999, consistent with the timing of the merger of Amoco and BP.[12] We also drop individual forecasters completely if the

12 See `bp.com/en/global/corporate/about-bp/our-history/heritage-brands.html`.

resulting panel is too short or incomplete, based on conditions that we outline later, or if there is a considerable break with no forecasts.[13]

A full set of the 33 US forecasters we use in our estimation is given in table A1, with the names that were dropped from the estimation listed in table A3. To point out one peculiarity in the dataset, "Wells Fargo" appears twice in the table, as the name of Forecaster 3 that emerged from the merger between Wachovia Corporation by Wells Fargo and Company on December 31, 2008,[14] and also as the name of Forecaster 33, sandwiched between two entries of "Wells Fargo Bank" before the forecaster name was changed to "Wells Capital."

We repeat the same process with Canadian forecasters, ending up with a panel of 22, as listed in table A2 (with the dropped names again listed in table A3). As with "Wells Fargo" for the US, "National Bank of Canada" appears twice in the Canadian panel: early in the sample (Forecaster 15) and later, after a 10-year absence, as the descendant of "National Bank Financial," which was itself formed from the merger of "First Marathon" and "Lévesque Beaubien Geoffrion Inc." (Forecaster 13).[15]

13 For example, Lehman Brothers was present in the panel twice: 03/1991–02/1994 and 12/2001–09/2008. We treat these as two separate forecasters, given the eight-year gap between them. We then drop the first period, as it is too short, but include the second, combined with Barclays Capital who took over Lehman's North American operations in September 2008 (`theguardian.com/business/2008/sep/16/barclay.lehmanbrothers1`), as Forecaster 16 in our panel.

14 See `wellsfargo.com/about/corporate/wachovia`.

15 See `nbc.ca/en/about-us/our-organization/the-bank/portrait-of-the-bank/history.html`.

TABLE A1
US forecasters

1	Amoco Corporation 10/1989–02/1999	BP Amoco 03/1999–04/1999			
2	Chrysler 04/1994–11/1998	Daimler Chrysler 12/1998–11/2007	Chrysler 12/2007–11/2008		
3	Core States 10/1989–05/1995	CoreStates Fin Corp. 07/1995–04/1998	First Union Corp. 06/1998–10/2001	Wachovia Corp. 11/2001–06/2009	Wells Fargo(1) 07/2009–12/2013
4	CRT Govt. Securities 09/1990–07/1993	NationsBank 08/1993–08/1998	Bank America Corp 11/1998–02/2009		
5	DuPont 10/1992–12/2013				
6	Eaton Corporation 11/1991–12/2013				
7	Econ Intelligence Unit 11/2003–12/2013				
8	Fannie Mae 10/1993–12/2013				
9	First Boston 10/1989–10/1994	CS First Boston 11/1994–04/1998	Credit Suisse First Boston 05/1998–10/2004	Credit Suisse 01/2009–12/2013	
10	Ford Motor 10/1989–12/2013				
11	General Motors 10/1989–12/2013				
12	Georgia State University 02/1996–12/2013				
13	Goldman Sachs 02/1999–12/2013				
14	Griggs & Santow 11/1989–09/2001				
15	Inforum – Univ. of Maryland 04/1998–12/2013				

continued

TABLE A1
continued

16	Lehman Brothers 12/2001–09/2008	Barclays Capital 01/2009–12/2013		
17	Macroeconomic Advisers 03/2000–12/2013			
18	Merrill Lynch 10/1989–04/2009	Bank of America – Merrill 05/2009–12/2013		
19	Morgan Guaranty 10/1989–07/1993	J P Morgan 08/1993–12/2013		
20	Morgan Stanley 10/1989–12/2013			
21	Mortgage Bankers Assoc. 10/1993–05/2005			
22	Natl. Assoc. of Home Builders 10/1993–12/2013			
23	Northern Trust 10/1989–12/2013			
24	Oxford Economics 10/1997–12/2013			
25	Prudential Insurance 12/1993–04/2001	Prudential Financial 05/2001–10/2002		
26	Regional Financial Assocs. 05/1994–01/2001	Economy.com 02/2001–11/2005	Moody's Economy.com 12/2005–05/2010	Moody's Analytics 06/2010–12/2013
27	Smith Barney 10/1989–06/1993	Smith Barney Shearson 08/1993–06/1994	Smith Barney 07/1994–10/1997	
28	Standard & Poor's 07/1990–12/2013			

continued

TABLE A1
continued

29	The Conference Board 10/1993–12/2013				
30	The University of Michigan 10/1993	Univ. of Michigan – RSQE 11/1993–12/2013			
31	The WEFA Group 07/1991–10/2001	DRI–WEFA 11/2001–11/2002	Global Insight 12/2002–10/2008	IHS Global Insight 11/2008–12/2013	
32	U.S. Trust 10/1993	United States Trust 11/1993–09/2007			
33	Wells Fargo Bank 01/1993–09/1994	Wells Fargo(2) 11/1994–02/1996	Wells Fargo Bank 04/1996–01/2000	Wells Capital 02/2000–06/2009	Wells Capital Mgmt. 07/2009–12/2013

TABLE A2
Canadian forecasters

1	Bank of Montreal 10/1989–01/2007			
2	Bank of Nova Scotia 10/1989–05/1998	Scotia Economics 06/1998–12/2013		
3	Caisse de dépôt 10/1989–06/2012	Caisse de dépôt		
4	CIBC 10/1989–08/1999	Canadian Imperial Bank 09/1999–02/2002		
5	Conf. Board of Canada, Conference Board 10/1989–12/2013			
6	Desjardins 08/2004–12/2013			
7	Economap 11/1999–12/2013			
8	EDC Economics 07/2003–12/2013			
9	Global Insight 12/2002–10/2008	IHS Global Insight 11/2008–12/2013		
10	Informetrica 11/1992–12/2013			
11	Institute of Policy Analysis 10/1995–11/1995	University of Toronto 12/1995–12/2013		
12	JP Morgan Canada 07/1995–08/1996	JP Morgan 11/1996–06/2008		
13	Levesque Beaubien 11/1993–08/1999	National Bank Financial 09/1999–06/2012	National Bank of Canada (1) 07/2012–12/2013	
14	Merrill Lynch Canada 11/2000–03/2012			
15	National Bank of Canada(2) 10/1989–04/2002			
16	Nesbitt Thomson 10/1989–08/1994	Nesbitt Burns 09/1994–02/2000	BMO Nesbitt Burns 03/2000–06/2006	BMO Capital Markets 07/2006–12/2013
17	RBC Dominion Securities, RBC Dominion, RBC – Dominion Securities 10/1989–04/2000			
18	Royal Bank of Canada 10/1989–12/2013			
19	Scotia McLeod 10/1989–03/1998			
20	Sun Life 10/1989–11/1998			
21	Toronto Dominion Bank 04/1999–12/2013			
22	Wood Gundy 10/1989–10/1995	CIBC Wood Gundy 11/1995–08/1999	CIBC Markets s 09/1999–12/1999	CIBC World Markets 01/2000–12/2013

TABLE A3
Dropped forecasters

US				
Action Economics	Chase Manhattan Bank	First Trust Advisors	Mellon Bank	Roubini Global Econ
American International Group	Chemical Bank	HSBC	Nat. Assn. of Manufacturers	Sears Roebuck
Bank of Boston	Chemical Banking	Kemper Financial	Nomura	Shawmut Bank
Bank One Corp	Citigroup	Lehman Brothers, 03/91–02/94	Paine Webber	Shawmut National
Bankers Trust	Continental Bank	Manufacturers Hanover	PNC Bank	Shearson Lehman
Bear Stearns	Dun & Bradstreet	Marine Midland	PNC Financial Services	Swiss Re
Bethlehem Steel	First Chicago	Mass Financial Services	Provident Bank	UBS
Brown Brothers	First Fidelity	Metropolitan Life	RDQ Economics	US Chamber of Commerce
Brown Brothers Harriman				
Canada				
Bank of America – Merrill	Centre for Spatial Economics	Econ Intelligence Unit	Merrill Lynch – Canada	Toronto Dominion, 10/89–07/92
Burns Fry	Citigroup	HSBC	Oxford Economics	UBS
Bunting Warburg	DRI Canada	Loewen Ondaatje	Richardson Greenshields	WEFA Canada
Capital Economics	Du Pont	McLean McCarthy	Royal Trust	

References

Autrup, S. L., and M. Grothe (2014) "Economic surprises and inflation expectations: Has anchoring of expectations survived the crisis?," ECB Working Paper Series, no. 1671

Bataa, E., D. Osborn, M. Sensier, and D. Dijk (2014) "Identifying changes in mean, seasonality, persistence and volatility for G7 and euro area inflation," *Oxford Bulletin of Economics and Statistics* 76(3), 360–88

Capistrán, C., and M. Ramos-Francia (2010) "Does inflation targeting affect the dispersion of inflation expectations?," *Journal of Money, Credit and Banking* 42(1), 113–34

Cecchetti, S. G., and C. Hakkio (2009) "Inflation targeting and private sector forecasts," NBER working paper no. 15424

Chowdhury, I., M. Hoffmann, and A. Schabert (2006) "Inflation dynamics and the cost channel of monetary transmission," *European Economic Review* 50(4), 995–1016

Clark, T. E., and T. Nakata (2008) "Has the behavior of inflation and long-term inflation expectations changed?," *Federal Reserve Bank of Kansas City Economic Review* Q1, 17–50

Davis, J. (2014) "Inflation targeting and the anchoring of inflation expectations: Cross-country evidence from consensus forecasts," Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute working paper no. 174

Detmeister, A., D. Jorento, E. Massaro, and E. Peneva (2015) "Did the Fed's announcement of an inflation objective influence expectations?," FEDS Notes 2015-06-08, Board of Governors of the Federal Reserve

Dovern, J., U. Fritsche, P. Loungani, and N. Tamirisa (2015) "Information rigidities: Comparing average and individual forecasts for a large international panel," *International Journal of Forecasting* 31(1), 144–54

Ehrmann, M. (2015) "Targeting inflation from below: How do inflation expectations behave?," *International Journal of Central Banking* 11(S1), 213–49

Galati, G., S. Poelhekke, and C. Zhou (2011) "Did the crisis affect inflation expectations?," *International Journal of Central Banking* 7(1), 167–207

Hofmann, B., and F. Zhu (2013) "Central bank asset purchases and inflation expectations," *BIS Quarterly Review* (March), 23–35

Johnson, D. R. (2002) "The effect of inflation targeting on the behavior of expected inflation: Evidence from an 11 country panel," *Journal of Monetary Economics* 49(8), 1493–519

Levin, A. T., F. B. Natalucci, and J. M. Piger (2004) "Explicit inflation objectives and macroeconomic outcomes," ECB Working Papers, no. 383

Mehrotra, A., and J. Yetman (2014) "Decaying expectations: What inflation forecasts tell us about the anchoring of inflation expectations," BIS Working Papers, no. 464

Morley, J., J. Piger, and R. Rasche (2015) "Inflation in the G7: Mind the gap(s)?," *Macroeconomic Dynamics* 19(04), 883–912

Nautz, D., and T. Strohsal (2015) "Are US inflation expectations re-anchored?," *Economics Letters* 127, 6–9

Shanno, D. F. (1985) "On Broyden–Fletcher–Goldfarb–Shanno method," *Journal of Optimization Theory and Applications* 46(1), 87–94

Siklos, P. L. (2013) "Sources of disagreement in inflation forecasts: An international empirical investigation," *Journal of International Economics* 90(1), 218–31

Stock, J., and M. Watson (2007) "Why has inflation become harder to forecast?," *Journal of Money, Credit and Banking* 39(s1), 3–33

Strohsal, T., and L. Winkelmann (2015) "Assessing the anchoring of inflation expectations," *Journal of International Money and Finance* 50, 33–48
Yetman, J. (2015) "The evolution of inflation expectations in Canada and the US," BIS Working Papers, no. 523