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INFLATION TARGETING AND PRIVATE SECTOR FORECASTS

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ABSTRACT

Transparency is one of the biggest innovations in central bank policy of the past quarter century. Modern central bankers believe that they should be as clear about their objectives and actions as possible. However, is greater transparency always beneficial? Recent work suggests that when private agents have diverse sources of information, public information can cause them to overreact to the signals from the central bank, leading the economy to be too sensitive to common forecast errors. Greater transparency could be destabilizing. While this theoretical result has clear intuitive appeal, it turns on a combination of assumptions and conditions, so it remains to be established that it is of empirical relevance.

In this paper we study the degree to which increased information about monetary policy might lead to individuals coordinating their forecasts. Specifically, we estimate a series of simple models to measure the impact of inflation targeting on the dispersion of private sector forecasts of inflation. Using a panel data set that includes 15 countries over 20 years we find no convincing evidence that adopting an inflation targeting regime leads to a reduction in the dispersion of private sector forecasts of inflation. While for some specifications adoption of inflation target does seem to reduce the standard deviation of inflation forecasts, the impact is rarely precise and always small.

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1 Introduction

Transparency is one of the biggest innovations in central bank policy of the past quarter century. Modern central bankers believe that they should be as clear about their objectives and actions as possible. This notion arises from the view that policymakers should be a source of stability, not a source of noise; with the economy and markets responding to data, not to the policymakers themselves.

Inflation targeting is one of the first and most comprehensive implementations derived from this view. As a framework for monetary policy, inflation targeting involves 'the public announcement of medium-term numerical targets for inflation [and] increased transparency of the monetary policy strategy through communication with the public and the markets about plans, objectives, and decisions of the monetary authority' (Mishkin 2002, pg. 361). The result is not only clearly understood and published numerical targets, but also inflation reports that explain past and likely future actions. Most economists believe that greater transparency is beneficial. See, for example, the survey papers by Walsh (2007), Carpenter (2004), Dincer and Eichengreen (2002), and Geraats (2002). However, transparency is not nudity. Understanding policymakers' contingency plans does not mean laying the policymaking process bare for all to see. Monetary policymakers should not put cameras in the meeting room. There are clear limits. What are they?

Recent theoretical work has put this question into a new perspective. In their pioneering work, Morris and Shin (2002 and 2005) show that when private agents have diverse sources of information, public information can cause them to overreact to the signals from the central bank, which makes the economy too sensitive to common forecast errors. The reason for this is that individuals care not only about accurately estimating the state of the economy, but also about having an estimate that is not too different from that of others. The implication is that more transparency may in fact be destabilising, so policymakers should think long and hard before changing their disclosure policies in ways that publicise more information.

Svensson (2006) and Woodford (2005) both suggest that the Morris and Shin result is likely to be a theoretical curiosum rather than anything policymakers should worry about. That is, the circumstance under which additional information is welfare reducing is extremely unlikely to occur in the real world. As Svensson shows, Morris and Shin's own conclusion only holds when the noise in policymakers' publicly announced information is at least eight times that of the private information agents have obtained on their own. That is, public officials must be far worse in their evaluations of the economic environment than private agents. Evidence, such as that in Romer and Romer (2000), suggests that central bank staff forecasts are at least as good, if not better than, those of market economists.

Woodford's critique is based on the Morris and Shin choice of how to aggregate the quadratic loss functions of the individual agents. In their original paper, Morris and Shin assume that policymakers seek to minimise a social loss function that is based on the average squared error of individual estimates of the state of the economy. By contrast, if the social welfare function includes losses associated with the dispersion of agents' estimates of the state – something Morris and Shin assume the agents themselves care about – then more information is unambiguously a good thing.

Regardless of these two coherent and largely convincing criticisms, the Morris and Shin argument retains intuitive appeal. In particular, policymakers worry that releasing more information might cause private agents to coordinate expectations, leaving the economy more exposed to common shocks. In the end, however, we are left with an empirical question: does increased transparency lead to lower dispersion in private forecasts? If the answer is no, then there is little to worry about. However, if the answer is yes, we cannot necessarily conclude that greater transparency—in the form of adopting an inflation target—is harmful. The reason is that greater transparency about the fundamentals and long-run inflation objective should also lead to a smaller dispersion of inflation forecasts which is beneficial. So, a smaller dispersion of private forecasts could reflect the beneficial effects of greater transparency and not the harmful effects described by Morris and Shin.

In this paper we study the degree to which increased information about monetary policy might lead to a reduction in the dispersion of inflation forecasts. By combining information about whether a country targets inflation with the dispersion of private sector forecasts of inflation, we seek to understand how inflation targeting affects private sector behaviour. In particular, does inflation targeting lead to a smaller or possibly larger dispersion of private sector inflation forecasts? If it leads to a larger dispersion, then there is no need to worry about the harmful effects of greater transparency discussed by Morris and Shin. However, if inflation targeting leads to a smaller dispersion, then there is at least some evidence that increased information could be harmful because it leads individuals to coordinate their forecasts à la Morris and Shin or could be beneficial because of additional information about central bank objectives and fundamentals.

In order to examine this, we estimate a series of simple models designed to measure the impact of inflation targeting on the dispersion of private sector forecasts of inflation. Using a panel data set that includes 15 countries over 20 years we find no convincing evidence that adopting an inflation targeting regime leads to a reduction in the dispersion of private sector forecasts of inflation. While for some specifications adoption of an inflation target does seem to reduce the standard deviation of inflation forecasts, in others it does not. And the precision of the estimates is rarely very high. The bulk of our evidence does not support the view that a shift to inflation targeting has resulted in a significant decline in the cross-sectional standard deviation of inflation forecasts across survey respondents.

Before proceeding, it is useful to note that our work is distinct from, but related to, two earlier papers. First, Mankiw, Reis and Wolfers (2004) examine the dispersion of inflation expectations in survey data and find that inflation expectations have become more concentrated around the mean as the level of inflation has fallen. At first glance this may seem as if it is a result that is more positive than ours. But, given that Mankiw, Reis and Wolfers only study US data, it is not possible to disentangle the impact of disinflation from increased Federal Reserve transparency.

Levin, Natalucci and Piger (2004) investigate how well the mean of inflation expectations has been anchored, also from survey data. They provide evidence on how inflation targeting has changed the dynamics of inflation. Their results suggest that the adoption of an explicit inflation target reduces the correlation of long-run inflation expectations with short-run movements in inflation, largely eliminating the link between expectations and realised inflation. Furthermore, Levin, Natalucci and Piger find that the adoption of an inflation targeting framework lowers the persistence of inflation, so that inflation behaves more like a random walk.

The remainder of this paper is organised in five sections. Section 2 provides a description of the data we use. This is followed in section 3 with a simple statistical analysis, and in section 4 with the results of more sophisticated regressions. Section 5 discusses some possible extensions and provides evidence on the robustness of the results. Section 6 provides a conclusion.

2 Description of the data

We study the dispersion of monthly survey-based inflation expectations for a number of countries from October 1989 to April 2009. The data are collected by Consensus Economics. Each month the firm surveys a large cross-section of professional forecasters – currently more than 700 world-wide – asking each one for their current and next calendar years' predictions for growth, inflation, unemployment, and short- and long-term interest rates in the countries that they follow. For each month, for each variable, Consensus Forecasts reports the high, low, and median forecast, as well as the standard deviation of survey responses. While Consensus Forecasts supplies forecast information for more than 70 countries, we restrict ourselves to the following 15: Australia, Canada, the Euro area, France, Germany, Italy, Japan, the Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, the UK and the US. For many of the results, we ignore the Euro area because data are only available starting in December 2002.

Table 1: Dates for which inflation forecast data are available

	Inflation targeting regime	Non-inflation targeting regime
Australia	June 1993 to April 2009	November 1990 to May 1993
Canada	February 1991 to April 2009	October 1989 to January 1991
Euro Area		December 2002 to April 2009
France		October 1989 to April 2009
Germany		October 1989 to April 2009
Italy		October 1989 to April 2009
Japan		October 1989 to April 2009
Netherlands		January 1995 to April 2009
New Zealand	December 1994 to April 2009	
Norway	March 2001 to April 2009	June 1998 to February 2001
Spain	January 1995 to June 1998	July 1998 to April 2009
Sweden	January 1995 to April 2009	
Switzerland	January 2000 to April 2009	June 1998 to December 1999
United Kingdom	October 1992 to April 2009	October 1989 to September 1992
United States		October 1989 to April 2009

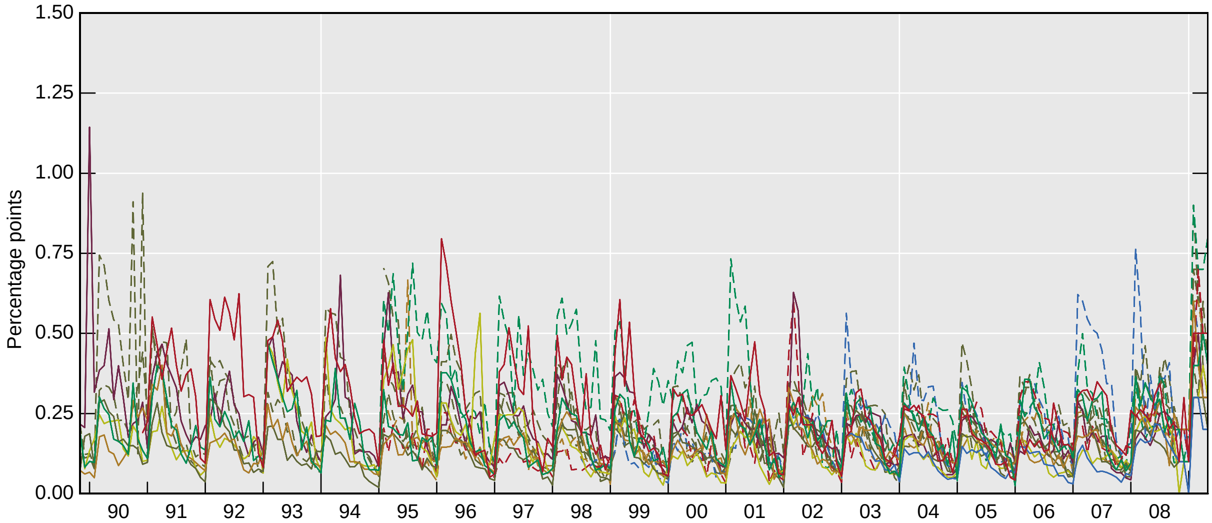
Notes: For the 'Inflation targeting regime' countries, the dates shown are either the first date of inflation targeting or the first date for which data are available; as a result, they do not necessarily correspond to dates at which a country adopted an inflation target. For the 'Non-inflation targeting regime' countries, the dates shown correspond to dates for which the dispersion of inflation and GDP forecasts are available.
Source: Appendix A, Mishkin and Schmidt-Hebbel (2007), Norges Bank Regulation on Monetary Policy, March 29, 2001. Contrary to the date given in Mishkin and Schmidt-Hebbel, the start date of inflation targeting for Australia is June 1993, based on data from the Reserve Bank of Australia.

This sample is sufficiently diverse to allow us to study the impact of inflation targeting, as two countries (New Zealand and Sweden) targeted inflation over the entire period, six (Australia, Canada, Norway, Spain, Switzerland, and the UK) adopted inflation targeting at some point during the sample, and the remaining eight have never adopted an explicit inflation target. For the second group, the six that adopted an inflation target during the 1990s, we need to choose a date for the adoption. It is perhaps surprising that there is disagreement on this timing. Mishkin and Schmidt-Hebbel (2007), Ball and Sheridan (2005), and Truman (2003), among others, all choose slightly different dates. For the most part, we adopt the dating in Appendix A of Mishkin and Schmidt-Hebbel (2007).

To continue, we need a bit of notation. We use the general form $S_{it}(.,.)$ to denote the standard deviation of private sector forecasts for country *i* made on date *t*. Next, we specify the variable being forecasted as π for CPI inflation and *y* for GDP growth, and whether the forecast is for the current year, which we denote by *c*, or for the next year, which we denote by n. Using this notation, $S_{it}(\pi, c)$ is the standard deviation of private sector forecasts for CPI inflation made at date t for the current year (the year containing t). Analogously, $S_{it}(\pi, n)$is the standard deviation of private sector forecasts for CPI inflation also made at date *t* but for the next year (the year containing t+1), and $S_{it}(y,c)$ and $S_{it}(y,n)$ are the standard deviations of private sector forecasts for GDP growth for the current and next year, respectively.

Figure 1: Standard deviation of current year's inflation forecasts, $S_{it}(\pi,c)$

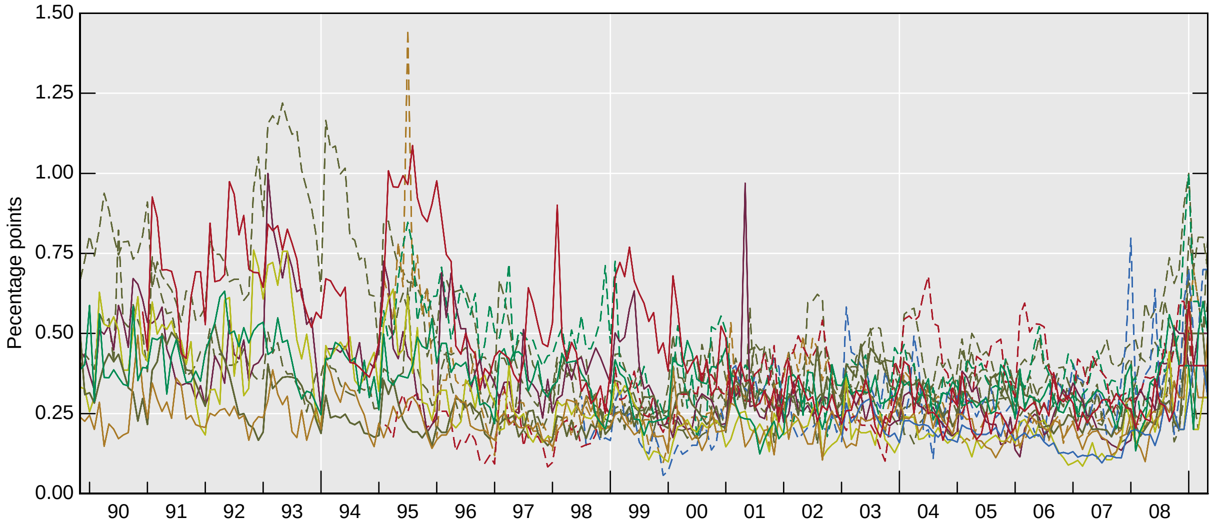
Monthly data



Source: Consensus Economics.

Figure 2: Standard deviation of next year's inflation forecasts, $S_{it}(\pi,n)$

Monthly data



Source: Consensus Economics.

To provide a sense of the time-series properties of the data, Figures 1 and 2 plot the standard deviation of current and next year's forecasts of inflation, $S_{it}(\pi,c)$ and $S_{it}(\pi,n)$, for all the countries in our sample. Simple inspection reveals that the standard deviation of forecasts for the current year, $S_{it}(\pi,c)$ plotted in Figure 1, has significant seasonality; while

$S_{it}(\pi,n)$ has less seasonality. Focusing on $S_{it}(\pi,c)$, a closer look shows that the standard deviation is highest in January and falls throughout the year. This is not at all surprising, since as a particular year progresses, inflation during that year increasingly becomes an historical fact that need not be estimated.

To assess the seasonality in these series, we assume they are deterministic and estimate the following regressions:

$$S_{it}(\pi,c) = \sum_{k=1}^{12} \beta_k^c D_{kt}^m + \varepsilon_{it} \tag{1}$$

and

$$S_{it}(\pi,n) = \sum_{k=1}^{12} \beta_k^n D_{kt}^m + \varepsilon_{it}\ , \tag{2}$$

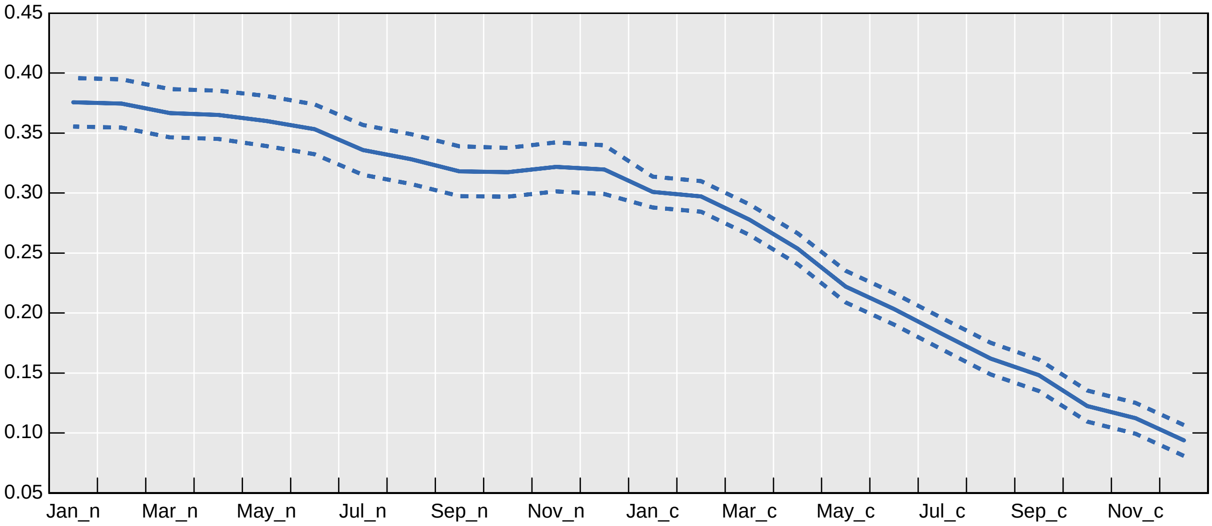
where D_{kt}^m is a monthly dummy for month k. Notice that for this exercise, the coefficients (the β_k's) are constrained to be equal across all of the countries in the sample.

The coefficients from equations (1) and (2), with a 95% confidence interval, are plotted in Figure 3. We report the coefficients from left to right depending on the amount of time from the date of the survey to the end of the period covered by the forecast. Starting on the far left, the first observation, labelled 'Jan_n,' is the coefficient associated with the dummy variable for the standard deviation of the January forecast for the next year's inflation (β_1^n), which is completed 24 months in the future. By contrast, the far right of the figure plots the coefficient associated with the standard deviation of inflation in December of the current year (β_{12}^c), which is completed in only one month. While the relationship is not linear, it is clearly declining. A regression of $\{\beta_1^n, \beta_2^n \ldots, \beta_{11}^c, \beta_{12}^c\}$ against {1, 2, ...,23, 24} yields a slope coefficient of -0.012 and a t-statistic of 16.2.[1] This means that the marginal effect of an additional month of data reduces the standard deviation of private inflation forecasts by 0.012. The average value of $S_{it}(\pi,n)$ in January (the standard deviation of private sector forecasts for 'next' year in January of 'this' year) is 0.377 percentage points and the average value of $S_{it}(\pi,c)$ in December is 0.094 percentage points. Thus, each additional month of data tightens the spread of private sector inflation forecasts (reduces the standard deviation) by 3.2 percent

[1] We also regressed the coefficients against {1, 2, …, 12, 0, …, 0} and {0, …, 0, 1, 2, …, 12}. The slope coefficients (t-statistics) were -0.003 (2.80) and -0.024 (23.1).

(0.012 / 0.377). The implication of all of this is that it is important that further analysis account for the pronounced seasonality in the data.

Figure 3: Seasonal dummy coefficients



Source: Author's calculations based on equations (1) and (2).

3 Simple statistical tests

We now turn to the key question of this paper: is the spread (standard deviation) of inflation forecasts by survey participants lower in countries that adopt an explicit inflation target? To examine this, we start with some very simple statistics designed to measure whether the spreads $S_{it}(\pi, c)$ and $S_{it}(\pi, n)$ are lower when a country's policymakers employ an inflation target. For clarity, we do this with a series of regressions. In the first one we estimate the following regression for each country separately:

$$S_{it}(\pi, c) = \alpha + \sum_{k=2}^{12} \beta_k D_{kt}^{m} + \gamma T \arg et_{it} + u_{it}^{c} \tag{3}$$

and similarly for $S_{it}(\pi, n)$. Of course, the variable 'Target' is only included if the country was both an inflation and non-inflation targeter during the sample period. In addition, since the available data are different for different countries, each equation is estimated with a different number of observations.

The results are shown in Appendix Tables A.1 and A.2 and summarised in Table 2. Note, there are only 6 countries that switched regimes during the sample period. For these countries, the average standard deviation for the non-inflation targeting regime is α, for the inflation targeting regime is $\alpha + \gamma$, and γ is the difference. If the dispersion is smaller for inflation targeting regimes, we would expect to see $\gamma < 0$.

Table 2: Comparing the standard deviation of inflation forecasts for countries that switch regime

γ	Standard deviation of current year's inflation forecast, $S_{it}(\pi, c)$	Standard deviation of next year's inflation forecast, $S_{it}(\pi, n)$
Negative, significant at 5%	Australia, UK	Australia, Canada, UK
Negative, insignificant at 5%	Spain, Switzerland	
Positive, insignificant at 5%	Canada	Switzerland
Positive, significant at 5%	Norway	Norway, Spain

Source: Authors' calculations based on country-by-country estimation of equation (3).

These results allow us to conclude that for those countries that adopted inflation targeting during the period 1990 to 2009, the standard deviation of private sector forecasts for CPI inflation sometimes falls and sometimes rises. Specifically, the standard deviation of inflation forecasts falls when Australia, Canada, and United Kingdom adopt inflation targets; but it increases when Norway adopts an inflation target. Spain and Switzerland are somewhere in between, depending on whether you are looking at the dispersion of current year forecasts or next year forecasts.

A next step is to estimate the equations jointly using Zellner's seemingly unrelated regression approach. Since the model is estimated as a set of equations, we need the same number of observations for all countries. There are basically three different start dates that could be used: November 1990, January 1995, or June 1998. Depending on the start date, the number of countries that were both inflation targeters and non-inflation targeters differs, as seen in Table 3. There is clearly a trade-off between sample size, number of countries, and mix of inflation regimes.

Table 3: Inflation regime status for various start dates

Start date	Switched	Non-targeter	Targeter only
November 1990 (T = 222, N = 8)	Australia, Canada, UK	France, Germany, Italy, Japan, US	
January 1995 (T = 172, N = 12)	Spain	France, Germany, Italy, Japan, Netherlands, US	Australia, Canada, New Zealand, Sweden, UK
June 1998 (T = 131, N =14)	Norway, Spain, Switzerland	France, Germany, Italy, Japan, Netherlands, US	Australia, Canada, New Zealand, Sweden, UK

Source: Table 1.

Table 4 summarises the results of estimation using Zellner's seemingly unrelated regression approach, focusing on the sign and significance of the inflation targeting variable for countries that were both inflation targeters and non-targeters. (Additional results are shown in Appendix Table A.3.) In general, the results are similar to the single equation regressions in Table 2: sometimes the coefficient is negative and sometimes it is positive; sometimes it is significantly different from zero at standard levels of statistical significance and sometimes it is not.

Table 4: Comparing the impact of inflation targeting on the standard deviation of inflation forecasts for inflation targeting and non-inflation targeting central banks using SUR

γ	Standard deviation of current year's inflation forecast, $S_{it}(\pi, c)$	Standard deviation of next year's inflation forecast, $S_{it}(\pi, n)$
Negative, significant at 5%	Australia, UK (1990) Spain (1995)	Australia (1990)
Negative, insignificant at 5%	Spain (1998)	U.K. (1990) Spain (1998)
Positive, insignificant at 5%	Canada (1990) Switzerland (1998)	Canada (1990)
Positive, significant 5%	Norway (1998)	Spain (1995) Norway (1998) Switzerland (1998)

Note: The dates in parentheses denote the start dates for the estimation.
Source: Authors' calculations based on SUR estimation of equation (3).

So far, the determinants of the dispersion of private sector forecasts are deterministic variables: seasonal dummies and an inflation targeting dummy. There are no economic variables that may affect the dispersion of forecasts. For example, the dispersion of inflation forecasts may be greater when overall macroeconomic variability is greater. While we do not have a country-specific measure of macroeconomic variability, we use the dispersion of private sector forecasts of GDP growth. To avoid any simultaneity concerns, we actually use the lagged value of the spread of GDP forecasts, $S_{i,t-1}(y,c)$ and $S_{i,t-1}(y,n)$. In addition, to capture the fact that volatility may have changed over time reflecting the so-called 'Great Moderation' and then the more recent financial crisis, we include a year variable and two dummy variables to capture the recent financial crisis. The two financial crisis variables are Crisis1 = 1 for September 2007 – September 2008, and Crisis2 = 1 for October 2008 – April 2009. These two variables reflect the fact that the financial crisis is generally thought to have started in August 2007 and then intensified in September 2008. Since the survey respondents presumably did not recognise the start and intensification of the crisis until the next month, the Crisis variables are dated one month after the start and intensification. In addition, we allow the current inflation dispersion to depend on its own lagged value, $S_{it-1}(\pi,c)$ and $S_{iit-1}(\pi,n)$. In all, we estimate a set of equations of the following form:

$$S_{it}(\pi,c) = \alpha_i^c + \sum_{k=2}^{12} \beta_k^c D_{kt}^m + \gamma^c Target_{it} + \delta^c year_t + \kappa_1^c Crisis_{1it} + \kappa_2^c Crisis_{2it} + \rho_i^c S_{it-1}(\pi,c) + \theta_i^c S_{it-1}(y,c) + u_{it}^c . \qquad (4)$$

Since the model is estimated as a set of equations, a *Target* variable can only be included for those countries that switched during the same period.

The results of estimating these equations using Zellner's seemingly unrelated regression approach are summarised in Table 5. With three different start dates, the countries that were both inflation targeting and non-inflation targeting differ. We report results only for the countries that change regime, so the coefficient of interest on *Target* can be estimated.

Table 5: Seemingly unrelated regression estimation

	Current year spread			Next year spread		
	Target	$S_{i,t-1}(\pi,c)$	$S_{i,t-1}(y,c)$	**Target**	$S_{i,t-1}(\pi,n)$	$S_{i,t-1}(y,n)$
January 1990 – April 2009						
Australia	-0.004 [0.2]	0.460 [7.7]	0.018 [0.3]	0.008 [0.3]	0.745 [16.9]	0.020 [0.5]
Canada	-0.006 [0.2]	0.409 [6.8]	-0.032 [0.7]	-0.008 [0.2]	0.573 [10.7]	-0.033 [0.9]
UK	0.009 [.4]	0.407 [8.5]	0.049 [0.8]	0.029 [1.4]	0.811 [23.3]	0.075 [1.6]
November 1995 – April 2009						
Spain	-0.016 [1.0]	0.344 [5.2]	-0.008 [0.1]	-0.014 [0.5]	0.637 [11.6]	0.121 [1.3]
June 1998 – April 2009						
Norway	0.034 [1.4]	0.432 [5.9]	0.-0.085 [1.7]	0.044 [1.5]	0.395 [5.5]	0.035 [0.7]
Spain	--- ---	0.521 [7.9]	-0.072 [1.1]	--- ---	0.635 [9.8]	-0.009 [0.1]
Switzerland	0.008 [0.5]	0.483 [7.3]	-0.013 [0.3]	0.029 [1.6]	0.322 [4.3]	0.096 [1.6]

Note: Numbers in brackets are asymptotic t-ratios.
Source: Authors' calculations based on equation (4).

This exercise allows us to draw several conclusions. First, the coefficients on *Target* reported in Table 5 are generally insignificant. However, the estimates of γ^c in the equation for $S_{it}(\pi,c)$ are positive and significant for Norway, Spain, and Switzerland. Second, the coefficient on the lagged value of the spread of the inflation forecast (ρ^c and ρ^n) is positive and significant, with the estimated impact in the regression of the dispersion in next year's inflation forecast larger than for the current year's inflation forecast ($\rho^n > \rho^c$). And third, in looking at the full set of regressions not reported in the table, the coefficient on the lagged value of *S(y,.)* is generally insignificant. When the coefficient is significant, it is mostly positive, although there are a couple of cases when the coefficient is negative and significant.

While these results are interesting, they fail to utilise information from the countries that either never adopted an inflation target or did so prior to the beginning of our sample. We now turn to a more sophisticated analysis designed to account for seasonality, control for general macroeconomic volatility, and employ all of the data we have available.

4 Panel regressions

The various shortcomings mentioned at the end of the previous section can be addressed by estimating a set of equations using a panel regression approach. By estimating various regressions using both fixed effects and random effects (Baum (2006)), we show that there is little evidence that inflation targeting countries have a smaller dispersion of private sector inflation forecasts.

Fixed Effects (FE)

We start by estimating some fixed effects models given the general applicability of this approach. To capture the fact that volatility may have changed over time reflecting the Great Moderation and then the more recent financial crisis, several dummy variables are included in the equation. Specifically, year dummy variables are included (Year1991, …, Year2006), where, for example, Year1991 = 1 for 1991. We also define Year2007 = 1 for January 2007 – August 2007, Crisis1 = 1 for September 2007 – September 2008, and Crisis2 = 1 for October 2008 – April 2009. The results from estimating the following equation (or some variant of it) are reported in Table 6 and the time-varying constants (the constant and terms related to time) are shown in Figure 4.[2]

$$S_{it}(\pi,c) = \alpha^c + u_i^c + \sum_{k=2}^{12} \beta_{ik}^c D_{kt}^m + \sum_{t=1991}^{2007} \delta_t Year_t + \kappa_1 Crisis_{1it} + \kappa_2 Crisis_{2it} \qquad (5)$$

$$+ \gamma_i^c Target_{it} + \rho^c S_{it-1}(\pi,n) + \theta^c S_{it-1}(y,n) + \varepsilon_{it}^c$$

where the constant α can be thought of as a mixture of the average level in January, the base month, plus the average impact of October 1989 to December 1990, while the *u* is the country-specific fixed effect. In addition, the table reports the long-run effect of inflation targeting on the dispersion: $\gamma/(1-\rho)$.

In all cases, the coefficient on Target is negative, which suggests that countries with an inflation target have a smaller standard deviation of private sector inflation forecasts. However, the coefficient is significant only for the bare-bones regression (excluding lagged

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[2] The chart plots the constant term (α) and coefficient on variables related to January of each year: $Year_t$ (δ_t), $Crisis_{1it}$ (κ_1), and $Crisis_{2it}$ (κ_2). In this way, the chart plots the "time varying constant term." Specifically, the value plotted for the years 1991 – 2007 are = $\alpha + \delta_t$ for 2008 = $\alpha + \kappa_1$, and for 2009 = $\alpha + \kappa_2$.

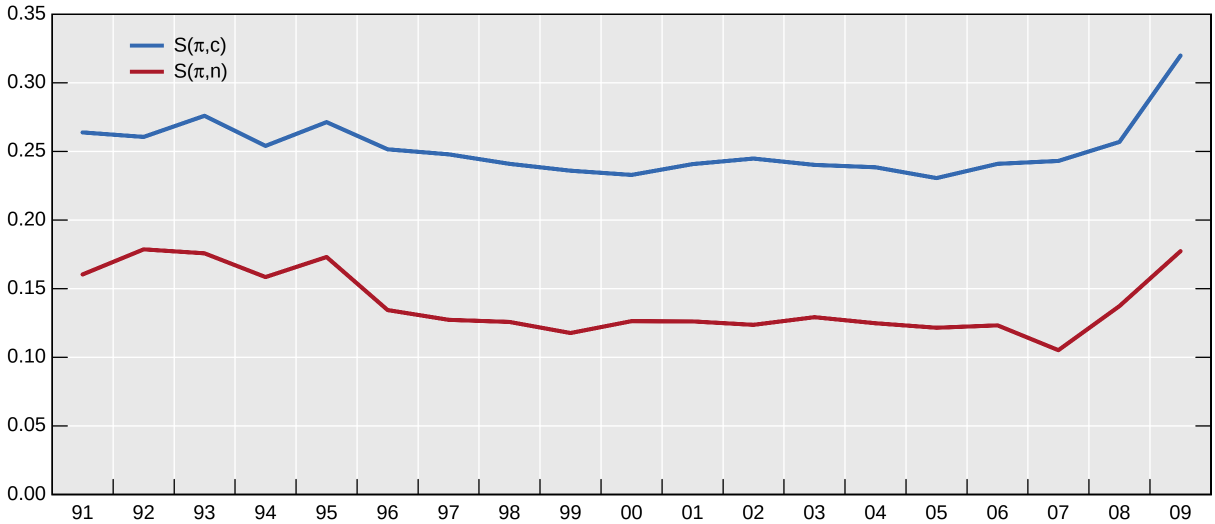
values of $S(\pi,.)$ and $S(y,.)$. In addition, the coefficients on lagged values of $S(\pi,.)$ and $S(y,.)$ are positive and significant. The estimate of ρ is 0.49 for regressions using $S_{it}(\pi,c)$ and 0.74 when looking at $S_{it}(\pi,n)$. This suggests that the persistence of the spread is less for current year forecasts than for next year forecasts, but still sizeable. A larger coefficient on the lagged spread for next year forecasts than for current year forecasts might suggest that incoming monthly data plays a smaller role for next year forecasts than for current year forecasts. In addition, to the extent that lagged values of $S(y,.)$ capture overall macroeconomic uncertainty, the positive and significant coefficients suggest that the standard deviation of inflation forecasts depend on macroeconomic uncertainty.

Table 6: Panel estimation using fixed effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi,c)$			Standard deviation of next year's inflation forecasts, $S_{it}(\pi,n)$		
	(1)	(2)	(3)	(4)	(5)	(6)
Inflation Target (γ)	-0.023	-0.011	-0.009	-0.021	-0.017	-0.009
	[3.0]	[1.4]	[1.3]	[2.0]	[1.6]	[1.3]
Lagged $S_{it}(\pi,\cdot)(\rho)$			0.488			0.737
			[29.1]			[57.3]
Lagged $S_{it}(y,\cdot)(\theta)$		0.142	0.036		0.103	0.015
		[9.1]	[2.6]		[6.1]	[1.3]
$\frac{\gamma}{(1-\rho)}$			-0.017			-0.036
			[1.3]			[1.3]
R^2	0.36	0.44	0.62	0.24	0.29	0.75
σ_ε	0.084	0.083	0.073	0.119	0.118	0.180
N	2895	2880	2880	2895	2880	2880

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly and yearly dummy variables are not shown.
Source: Authors' calculations based on equation (5).

Figure 4: Time-varying constant term



Source: Author's calculations.

Figure 4 shows the effect of time and the financial crisis on the standard deviation of private sector forecasts of current and next year inflation made in January. In general, the standard deviation declined from 1991 to 1999, was stable through 2006, and then rose significantly in 2008 and 2009. This suggests that the 'Great Moderation' did lead to a reduction in the dispersion of private sector forecasts of inflation through the first part of the sample period and that the financial crisis led to an increase in the dispersion. Interestingly, the coefficient on (α + δ) in 2009 is about the same as in the early 1990s for $S(\pi, n)$ but higher for $S(\pi, c)$.

While the coefficient on Target is negative for both $S(\pi, c)$ and $S(\pi, n)$, the magnitude is economically small. There are several ways to see this. First, look back at Figures 1 and 2 and note that the average of $S(\pi, c)$ is roughly 0.20 while $S(\pi, n)$ averages closer to 0.35. This means that the estimated impact of inflation targeting even in the long run (0.017 and 0.036) is to reduce the standard deviation of inflation forecasts by 10 percent or less.

A second way to see that γ is economically small is to compare it to the impact of the financial crisis. To do this, we consider what would happen to $S(\pi, c)$ if a country were to adopt an inflation target in February 2007 versus adopting it in February 2009. Taking the case of the United States, we use equation (2) to calculate the predicted value of $S(\pi, c)$ first with the *Target* = 0 (no inflation target) and then with the *Target* = 1 (assuming the US had an inflation target). The result is that the predicted dispersion in survey inflation expectations would fall from 0.269 to 0.260 in February 2007, but from 0.663 to 0.654 in February 2009. In other words, while dispersion would be less, the impact of financial turmoil on the dispersion is much larger –0.394 versus –0.036, or 11 times larger.

Random Effects (RE)

We next considered estimating the panel regression using a random effects estimator. The fixed effects (FE) model specifies the country specific effect as a constant, whereas the random effects (RE) model specifies the country specific effect as a random variable that is uncorrelated with the regressors. Breusch and Pagan (1980) developed a Lagrange multiplier test for $\sigma_u^2 = 0$; the p-value is reported in Table 9 in the row labeled '$\sigma_u^2 = 0$ p-value.' If the orthogonality assumption is true, then the random effects model is more efficient because it uses the assumption that u_i is uncorrelated with the regressors. Of course, if this assumption is false, then the random effects model is inconsistent. We can then use a Hausman test of the extra orthogonality condition imposed by the random effects estimator. The idea of the Hausman test is simple: if the regressors are uncorrelated with u_i, the fixed effects estimator is consistent but inefficient and the random effects estimator is consistent and efficient; however, if the regressors are correlated with u_i, the fixed effects estimator is consistent but the random effects estimator is inconsistent. Table 7 estimates the same models and includes the p-value from the Hausman test in the row labelled 'Hausman'.[3]

Table 7: Panel estimation using random effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi, c)$			**Standard deviation of next year's inflation forecasts,** $S_{it}(\pi, n)$		
Inflation Target (γ)	-0.004 [0.6]	0.013 [2.1]	0.022 [7.2]	-0.012 [1.2]	-0.007 [0.7]	0.014 [4.2]
Lagged $S_{it}(\pi,\cdot)$ (ρ)			0.542 [33.2]			0.785 [66.3]
Lagged $S_{it}(y,\cdot)$ (θ)		0.168 [11.2]	0.082 [7.0]		0.110 [6.5]	0.040 [4.2]
$\frac{\gamma}{(1-\rho)}$			0.047 [7.4]			0.066 [4.4]
R^2	0.39	0.49	0.65	0.25	0.31	0.76
σ_ε	0.084	0.083	0.073	0.119	0.118	0.080
$\sigma_u^2 = 0$ p-value	0.000	0.000	0.000	0.000	0.000	0.000
Hausman p-value	0.276	0.000	0.000	0.961	0.999	0.000
N	2895	2880	2880	2895	2880	2880

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly dummy variables are not shown.
Source: Author's calculations based on equation (5).

The results from using random effects to estimate the model are mixed. The coefficient on *Target* is sometimes negative and insignificant and other times it is positive and significant. Not surprisingly, one can always reject the hypothesis that $\sigma_u^2 = 0$. Unfortunately, one can often reject the hypothesis that the orthogonality condition holds. In general, it appears as though we cannot reject the hypothesis that the orthogonality condition holds for the stripped-down model (which includes only *Target* and year and month dummies) but we can generally reject the orthogonality condition when we include lagged values of $S(\pi, .)$ and $S(y, .)$. Interestingly, in the bare-bones model (which fails to reject the Hausman test) the coefficient on *Target* is negative but insignificant.

[3] Including a lagged dependent variable in a fixed effects model creates a large-sample bias in the estimate of the coefficient on lagged dependent variable. Since we are not really interested in this estimate, the concern is somewhat mitigated. In addition, in a simple model Nickell (1981) shows that for large values of T, the limit of $(\hat{\rho} - \rho)$ as $N \to \infty$ is approximately –(1+ρ)/(T-1). With ρ = 0.5 (0.8) and T = 235, the bias will be -0.0064 (-0.0077).

5 Robustness and Extensions

In this section, we check the robustness of the results by considering six modifications to the baseline model:

- we examine the sensitivity of the results to individual countries
- we compute the standard errors of the estimated coefficients using alternative techniques
- we consider the effect of the dating of the introduction of inflation targeting
- we include various measures of actual inflation
- we introduce commodity prices into the model
- we replace the standard deviation of forecasts with their root mean square error.

In what follows, we discuss results only for the fixed effects panel regression using $S_{it}(\pi, c)$ and $S_{it}(\pi, n)$. Thus, each of the six modifications to the baseline model is compared to the results in Table 6. In general, the conclusions from the previous section are supported.

Sensitivity to individual countries

Since the panel includes 15 countries, the first robustness check is to see whether some countries are 'influential'. To check this, the fixed-effects panel regression is estimated with all countries, and then we exclude one country at a time. So, for example, we re-estimate the model with Australia excluded, we then include Australia but exclude Canada, we then include Australia and Canada, but exclude Norway, and continue in this way.

Table 8 reports the coefficients and t-statistics on Target (γ) for all three panel equations for all countries, corresponding to the columns as labelled in Table 6, and then for each excluded country. In looking at the table, the appropriate comparison is between the first row ('None' excluded) and each subsequent row.

The results suggest that Canada and Norway may be influential for $S(\pi, c)$ and that Canada, Italy, the Netherlands, Norway and Spain, may be influential for $S(\pi, n)$. First, note that in all but 6 cases – 2 for $S(\pi, c)$ and 4 for $S(\pi, n)$ – the sign remains negative. In the 6 cases when the sign changes (Australia and the United Kingdom), the coefficient is small and insignificant. More importantly, Canada and Norway appear to be influential for $S(\pi, c)$ because the coefficient is insignificant in columns 2 and 3 when all countries are included, but becomes significant when Canada and Norway are excluded. This suggests that Canada and Norway are the countries behind the insignificant coefficient when all countries are included. These result may not be too surprising since the coefficient on Target was positive and insignificant for Canada and positive and significant for Norway in Tables 2, 4, and 5. In addition, when looking at the results for $S(\pi,\ n)$, Canada, Italy, the Netherlands, Norway and Spain are influential in the same way: the coefficient on Target becomes significant when these countries are excluded in columns 5 or 6. As with $S(\pi,\ c)$, the earlier results in Tables 2, 4, and 5, these results are not surprising given.

Table 8: Are Countries Influential?

Fixed-effects panel estimation of impact of inflation targeting (γ^c and γ^n)

Excluded Country	Standard deviation of current year's inflation forecasts, $S_{it}(\pi, c)$			Standard deviation of next year's inflation forecasts, $S_{it}(\pi, n)$		
	(1)	(2)	(3)	(4)	(5)	(6)
None	-0.023	-0.011	-0.009	-0.021	-0.017	-0.009
	[2.98]	[1.36]	[1.30]	[1.98]	[1.59]	[1.28]
Australia	-0.007	0.003	-0.000	0.008	0.011	-0.001
	[0.80]	[0.41]	[0.06]	[0.68]	[0.94]	[0.08]
Canada	-0.033	-0.020	-0.014	-0.030	-0.025	-0.012
	[4.06]	[2.40]	[1.95]	[2.62]	[2.17]	[1.48]
Norway	-0.058	-0.045	-0.028	-0.053	-0.050	-0.019
	[6.96]	[5.28]	[3.77]	[4.45]	[4.14]	[2.28]
Spain	-0.014	-0.001	-0.003	-0.039	-0.035	-0.013
	[1.60]	[0.09]	[0.41]	[3.14]	[2.76]	[1.51]
Switzerland	-0.022	-0.008	-0.008	-0.026	-0.020	-0.010
	[2.69]	[0.93]	[1.08]	[2.22]	[1.73]	[1.28]
United Kingdom	-0.005	0.005	-0.000	0.006	0.011	-0.003
	[0.59]	[0.64]	[0.04]	[0.54]	[0.99]	[0.35]
Euro area	-0.023	-0.011	-0.009	-0.021	-0.017	-0.010
	[2.96]	[1.34]	[1.27]	[1.96]	[1.58]	[1.27]
France	-0.020	-0.008	-0.007	-0.014	-0.010	-0.008
	[2.52]	[1.02]	[1.05]	[1.22]	[-0.91]	[1.00]
Germany	-0.022	-0.009	-0.008	-0.019	-0.014	-0.008
	[2.72]	[1.11]	[1.11]	[1.70]	[1.27]	[1.08]
Italy	-0.018	-0.006	-0.006	-0.027	-0.024	-0.011
	[2.30]	[0.70]	[0.84]	[2.45]	[2.10]	[1.49]
Japan	-0.028	-0.013	-0.010	-0.021	-0.016	-0.009
	[3.62]	[1.64]	[1.53]	[1.96]	[1.48]	[1.26]
Netherlands	-0.024	-0.013	-0.010	-0.024	-0.022	-0.011
	[3.13]	[1.60]	[1.42]	[2.28]	[2.04]	[1.44]
United States	-0.023	-0.011	-0.009	-0.016	-0.012	-0.008
	[2.91]	[1.36]	[1.28]	[1.48]	[1.13]	[1.11]
New Zealand	-0.023	-0.013	-0.011	-0.020	-0.017	-0.009
	[3.19]	[1.79]	[1.64]	[1.87]	[1.56]	[1.24]
Sweden	-0.021	-0.009	-0.008	-0.022	-0.017	-0.009
	[2.27]	[1.20]	[1.20]	[2.01]	[1.58]	[1.26]

Note: Asymptotic t-ratios are in brackets. The coefficient and t-statistic on inflation target (γ) are reported. Numbers in parentheses at the top of the columns refer to those in Table 6.

Table 9: Alternative Standard Errors

Panel Estimation using fixed effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi,c)$			Standard deviation of next year's inflation forecasts, $S_{it}(\pi,n)$		
Inflation Target (γ)	-0.023	-0.011	-0.009	-0.021	-0.017	-0.009
	[3.0]	[1.4]	[1.3]	[2.0]	[1.6]	[1.3]
	[0.6]	[0.3]	[0.5]	[0.5]	[0.4]	[0.8]
	[0.6]	[0.3]	[0.4]	[0.4]	[0.4]	[0.6]
Lagged $S_{it}(\pi,\cdot)(\rho)$			0.488			0.737
			[29.1]			[57.3]
			[10.0]			[17.3]
			[12.2]			[15.1]
Lagged $S_{it}(y,\cdot)$ (θ)		0.142	0.036		0.103	0.015
		[9.1]	[2.6]		[6.1]	[1.3]
		[3.8]	[1.9]		[2.0]	[1.3]
		[4.6]	[2.1]		[2.2]	[1.2]
$\frac{\gamma}{(1-\rho)}$			-0.017			-0.036
			[1.3]			[0.0]
			[0.5]			[0.8]
			[0.4]			[0.7]
R^2	0.36	0.44	0.62	0.24	0.29	0.75
σ_ε	0.084	0.083	0.073	0.119	0.118	0.080
N	2895	2880	2880	2895	2880	2880

Note: Three sets of t-statistics, all in brackets, are provided for each coefficient estimate. The first is the asymptotic t-statistic computed using the conventionally computed standard error. The second is based on a standard error that is robust to heteroskedasticity. And the third is computed using a bootstrap. Coefficients on monthly and yearly dummy variables are not shown.
Source: Authors' calculations based on equation (5).

Alternative measure of standard errors

Given the range of countries included in the panel of countries, the conventional standard errors reported in Table 6 may be misleading. Therefore, another robustness check involves estimating alternative standard errors. The standard errors reported in Table 6 are the typical standard errors for generalised least squares. As a check, Table 9 reports heteroskedasticity-robust standard errors and bootstrap standard errors.

Not surprisingly, the t-statistics are smaller. In particular, the t-statistic on the Inflation Target is now insignificant in all cases when using the robust or bootstrap standard errors. However, the other coefficients generally remain significant (the coefficient on lagged $S_{it}(y,n)$ is insignificant no matter what standard errors are used).

Dating the introduction of inflation targeting

As noted earlier, the dating of the inflation-targeting regime is somewhat ad hoc and different researchers use different dates for the start of the regime. In addition, if survey participants take time to learn about the inflation targeting regime (for example, how serious are the authorities?), then an alternative dating regime could give different results. To test this hypothesis, the fixed effects panel regressions are re-estimated using the third lag of Target rather than the current value of Target (from Table 1), thus allowing survey participants three months to learn about the new regime. The results are shown in Table 10.

In this case, the coefficients on the lagged inflation target variable are somewhat larger (in absolute value) and significant for all three models for $S_{it}(\pi, c)$, compared to only one of the models in Table 6. And the coefficient is significant for two out of three models for $S_{it}(\pi, n)$, compared to only one of the models in Table 6. Thus, there is some evidence that it takes a couple of months for survey participants to respond to the introduction of an inflation targeting regime. Stated somewhat differently, it means the first three months of inflation targeting are influential observations in models 2, 3, and 5 because the coefficient on the inflation targeting variable is insignificant when these observations are included but the coefficient is significant when these observations are excluded.

Table 10: Effect of Lagging the Date of Introduction of Inflation Targeting

Panel Estimation using fixed effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi, c)$			**Standard deviation of next year's inflation forecasts,** $S_{it}(\pi, n)$		
Lagged (3) Inflation Target (γ)	-0.029 [3.9]	-0.017 [2.3]	-0.014 [2.1]	-0.030 [2.7]	-0.026 [2.4]	-0.006 [0.8]
Lagged $S_{it}(\pi, \cdot)$ (ρ)			0.485 [29.6]			0.737 [57.2]
Lagged $S_{it}(y, \cdot)$ (θ)		0.137 [9.0]	0.033 [2.4]		0.101 [5.9]	0.012 [1.0]
$\frac{\gamma}{(1-\rho)}$			-0.027 [2.1]			-0.025 [0.9]
R^2	0.49	0.50	0.62	0.33	0.34	0.69
σ_ε						
N	2850	2850	2850	2850	2850	2850

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly and yearly dummy variables are not shown.
Source: Authors' calculations based on equation (5) with *Target* lagged 3 months.

Including measures of actual inflation

Turning now to some extensions of the baseline model, we first introduce various measures of actual inflation into the specification. One measure of inflation is the percent change in the CPI from 12 months ago and the other measure is the year-average-over-year-average percent change. Since the survey participants provide estimates of the year-average-over-year-average inflation rate, the percent change from 12 months ago is not exactly comparable to the inflation rate being forecasted. However, it does provide actual inflation data that survey participants could use in forecasting inflation. The year-average-over-year-average inflation rate is comparable, but is the same for all months in the year. Including the lagged value of either inflation rate does not change the results in any meaningful way: the t-statistic on the inflation rate is always less than 1.0. (To conserve space, the results are not shown.)

Another inflation measure turned out to be significant. Since the dependent variable is the dispersion of private sector forecasts, one might expect that the variability of inflation would be a significant explanatory variable. The results in Table 6 suggest that the dispersion of private sector forecasts is a significant explanatory variable. We extend those results by including a measure of the variability of actual inflation, measured as the 12-month rolling standard deviation of month-over-year-ago inflation rates. This variable is significant in all six cases; the t-statistic is between 6.5 and 12.3 in equations for $S(\pi, c)$ and between 4.1 and 10.1 in equations for $S(\pi, n)$. However, the coefficients and t-statistics on the other variables are fairly similar, so the results are not shown (to conserve space).

Table 11: Including Volatility of Commodity Price Inflation

Panel Estimation using fixed effects

	Standard deviation of current year's inflation forecasts, $S_{it}(\pi, c)$			**Standard deviation of next year's inflation forecasts,** $S_{it}(\pi, n)$		
Inflation Target (γ)	-0.014 [1.9]	-0.006 [0.8]	-0.004 [0.5]	0.003 [0.3]	0.005 [0.4]	-0.004 [0.5]
Lagged $S_{it}(\pi, \cdot)$ (ρ)			0.495 [29.7]			0.725 [54.2]
Lagged $S_{it}(y, \cdot)$ (θ)		0.127 [8.6]	0.032 [2.4]		0.092 [5.5]	0.017 [1.4]
$\frac{\gamma}{(1-\rho)}$			-0.007 [0.5]			-0.016 [0.5]
R^2	0.363	0.442	0.660	0.135	0.163	0.726
σ_ε	0.078	0.077	0.067	0.114	0.114	0.079
N	2811	2803	2803	2811	2803	2803

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly and yearly dummies are not shown.
Source: Authors' calculations based on equation (5) with commodity prices added.

Commodity prices

Given the wide range of countries used in the panel, the effect of having an inflation target may be overwhelmed by other factors. In particular, some countries are more susceptible to commodity price shocks than other countries. In an effort to check this, we extend the results in Table 6 by including a measure of the volatility of commodity price inflation. Specifically, we include the lagged value of the rolling 12-month standard deviation of the percent change (from the previous month) in a commodity price index; the index used is the Commodity Research Bureau spot raw industrial price index. The same commodity price index is used for all countries under the assumption that commodity prices are set in world markets. However, the coefficient on the commodity price variable is allowed to be different for each country. The results are shown in Table 11. Recall that there are 15 countries and three equations for $S_{it}(\pi, c)$, so there are 45 coefficients for commodity price variability. In order to conserve space the coefficients on commodity price variability are not shown. Briefly, the coefficients on commodity price variability are almost always positive and significant. For the $S_{it}(\pi, c)$ equations, 36 of the 45 coefficients are positive and significant, and the other nine coefficients are positive and insignificant. For the $S_{it}(\pi, n)$ equations, 35 of the coefficients are positive (24 are significant and 11 are insignificant) and 10 of the coefficients are negative (two are significant and eight are insignificant). Returning to our primary interest, however, we note that when we add the commodity price variation measure to the regression, the coefficient on Inflation Target is similar in magnitude and significance in the $S_{it}(\pi, c)$ equations, but about half the magnitude and significance in the $S_{it}(\pi, n)$ equations.

Root-mean-square error

Finally, we estimated a model similar to equation (3) but using the root-mean-squared-error of the inflation forecast, $RMSE_{it}(\pi, c)$ and $RMSE_{it}(\pi, n)$ rather than using the standard deviation of current (and next) year's inflation forecast. (Appendix 1 describes the mechanics of how the *RMSE* is recovered from the data we have available.) Specifically, we estimate the following equation by fixed-effects:

$$RMSE_{it}(\pi, n) = \alpha^n + u_i^n + \sum_{k=2}^{12} \beta_{ik}^n D_{kt}^m + \sum_{t=1991}^{2007} \delta_t^n Year_t + \kappa_1^n Crisis_{1it} + \kappa_2^n Crisis_{2it}$$

$$+ \gamma_i^n Target_{it} + \rho^n RMSE_{1t-1}(\pi, n) + \theta^n S_{1t-1}(y, n) + \varepsilon_{it}^n \quad (6)$$

Table 12 reports the results of this exercise. These are clearly more compelling than what we obtained using the dispersion of private sector forecasts against an inflation targeting variable. In particular, the coefficient on Inflation Target is negative and significant in 5 of the 6 equations.

While it may appear that the *RMSE* results are stronger than those using the dispersion, we view them with caution for two reasons. First, while the coefficient on the inflation target variable is larger when using RMSE than when using $S_{it}(\pi,.)$, the RMSE itself is also larger than $S_{it}(\pi,.)$. For example, the average (across all countries and time) of $S_{it}(\pi, c)$ is 0.20 while the average of $RMSE_{it}(\pi, c)$ is 0.40; and the average of $S_{it}(\pi, n)$ is 0.34 while the average of $RMSE_{it}(\pi, c)$ is 0.77. A better way to compare the results using $S_{it}(\pi,.)$ and

RMSE is to look at the effect of adopting an inflation target relative to the average value of $S_{it}(\pi,.)$ and $RMSE_{it}(\pi,.)$as we did above. Here the results are the same, the reduction in $S_{it}(\pi,.)$or $RMSE_{it}(\pi,.)$is small relative to the average size of S(π, .) or $RMSE_{it}(\pi,.)$for non-inflation targeting regimes. The reduction is between 3 percent and 47 percent, depending on which model is used and whether we look at current year or next year forecasts. Second, the $RMSE_{it}(\pi,.)$is much more volatile than $S_{it}(\pi,.)$. And not only is it more volatile, the volatility is more episodic in that there are some years when $RMSE_{it}(\pi,.)$is three to five times larger than $S_{it}(\pi,.)$and other years when they are about the same magnitude. Part of the reason is that $RMSE_{it}(\pi,.)$ is calculated from relative actual inflation measured as year-average-over-year-average percent change and is thus the same for all months in a given year.

Table 12: Using the Root-Mean-Square Error

Panel Estimation using fixed effects

	Standard deviation of current year's inflation forecasts, $RMSE_{it}(\pi,c)$			**Standard deviation of next year's inflation forecasts,** $RMSE_{it}(\pi,n)$		
Inflation Target (γ)	-0.163 [7.0]	-0.126 [5.3]	-0.019 [1.3]	-0.213 [5.9]	-0.209 [5.7]	-0.043 [2.3]
Lagged $RMSE_{it}(\pi,\cdot)$ (ρ)			0.790 [68.1]			0.859 [87.6]
Lagged $S_{it}(y,\cdot)$ (θ)		0.350 [6.9]	0.017 [0.5]		0.231 [4.0]	0.034 [1.1]
$\frac{\gamma}{(1-\rho)}$			-0.091 [1.3]			--0.303 [2.7]
R^2	0.10	0.16	0.79	0.07	0.10	0.84
σ_ε	0.257	0.254	0.156	0.396	0.395	0.203
N	2835	2820	2820	2820	2805	2805

Note: Asymptotic t-ratios are in brackets. Coefficients on monthly and yearly dummy variables are not shown.
Source: Authors' calculations based on equation (7) with the RMSE as the left-hand-side variable.

6 Conclusions

Using survey data on inflation expectations drawn from Consensus Forecast, we find little evidence that inflation targeting countries have a smaller dispersion of private sector forecasts of inflation. While for some countries, some models, and some estimation techniques, we estimate that inflation targeting countries have a smaller dispersion of private sector inflation forecasts, for other countries, other models, and other estimation techniques, we find that they do not.

Returning to the question that motivated this analysis – 'Does increased transparency lead to lower dispersion in private forecasts?' – the answer appears to be no. This suggests to us that the Morris and Shin argument that increased transparency could be destabilising is of little practical concern to policymakers. And even in the cases when inflation targeting leads to a lower dispersion of private forecasts, we cannot necessarily conclude that greater transparency is harmful because the lower dispersion could reflect the beneficial effects from greater transparency rather than the harmful effects from greater coordination. Of course, since the survey we use only reports results of forecasts for inflation in the current year and the next year, our results are unable to shed light on whether the distribution of private sector forecasts of long-run inflation is lower. However, even if the inflation target is for the medium-run, one would expect that the dispersion of inflation forecasts for ‘next year’ may still be somewhat smaller than otherwise.

Appendix

Table A.1: Average standard deviation of current year inflation forecasts

	α—January			
	No inflation targeting	**Inflation targeting**	**Difference**	**Number of observations**
Australia	0.513	0.377	-0.136	222
	[17.8]	[15.9]	[6.86]	
Canada	0.294	0.296	0.002	235
	[15.8]	[22.8]	[0.1]	
Norway	0.277	0.418	0.141	131
	[8.2]	[13.6]	[7.0]	
Spain	0.249	0.229	-0.021	172
	[11.8]	[9.8]	[1.4]	
Switzerland	0.313	0.298	-0.015	131
	[12.3]	[14.7]	[0.9]	
UK	0.473	0.322	-0.151	235
	[14.5]	[11.9]	[6.9]	
Euro Area	0.169			77
	[10.9]			
France	0.227			235
	[17.3]			
Germany	0.211			235
	[26.2]			
Italy	0.234			235
	[12.6]			
Japan	0.281			235
	[11.6]			
Netherlands	0.254			172
	[12.0]			
US	0.339			235
	[18.3]			
New Zealand		0.467		173
		[12.7]		
Sweden		0.349		172
		[15.0]		

Table A.2: Average standard deviation of next year inflation forecasts				
	α—January			
	No inflation targeting	**Inflation targeting**	**Difference**	**No inflation targeting**
Australia	0.791 [14.5]	0.496 [11.0]	-0.295 [7.8]	222
Canada	0.426 [13.8]	0.364 [16.9]	-0.062 [2.5]	235
Norway	0.226 [5.7]	0.354 [9.7]	0.127 [5.3]	131
Spain	0.315 [7.9]	0.416 [9.5]	0.101 [3.7]	172
Switzerland	0.269 [10.1]	0.300 [14.0]	0.031 [1.7]	131
UK	0.771 [12.0]	0.506 [9.5]	-0.265 [6.2]	235
Euro Area	0.231 [9.4]			77
France	0.245 [16.1]			235
Germany	0.318 [17.2]			235
Italy	0.304 [8.9]			235
Japan	0.403 [12.2]			235
Netherlands	0.338 [10.4]			172
US	0.452 [17.5]			235
New Zealand		0.442 [12.2]		173
Sweden		0.327 [11.0]		172

Table A.3: Coefficient on inflation targeting dummy variable				
		November 1990 – April 2009	**January 1995 – April 2009**	**June 1998 – April 2009**
Australia	$S_{it}(\pi,c)$	-0.109 [5.9]		
	$S_{it}(\pi,n)$	-0.182 [6.0]		
Canada	$S_{it}(\pi,c)$	0.049 [1.7]		
	$S_{it}(\pi,n)$	0.004 [0.1]		
UK	$S_{it}(\pi,c)$	-0.077 [4.3]		
	$S_{it}(\pi,n)$	-0.059 [1.9]		
Spain	$S_{it}(\pi,c)$		-0.029 [3.0]	-0.075 [1.6]
	$S_{it}(\pi,n)$		0.080 [4.5]	-0.015 [0.2]
Norway	$S_{it}(\pi,c)$			0.135 [8.3]
	$S_{it}(\pi,n)$			0.117 [6.8]
Switzerland	$S_{it}(\pi,c)$			0.008 [1.1]
	$S_{it}(\pi,n)$			0.040 [2.4]

Calculation of the root-mean-squared-error of the inflation forecast

Let $\pi_{j,t}^{fi,c}$ = inflation forecast (f) for country i, for the current year, made by forecaster j at time t and π_t^{ai} = actual inflation for country i at time t. Since the inflation forecasts are for year/year (rather than month/month-year-ago), π_t^{ai} is the same for each month in the year corresponding to date t: $\pi_{Jan1990}^{ai} = \pi_{Feb1990}^{ai} = \ldots = \pi_{Dec1990}^{ai}$. Next, let E^j denote taking expectations (average) over survey participants, j. Then, $E^j(\pi_{j,t}^{fi,c}) = \overline{\pi}_t^{fi,c}$ (which no longer has the j subscript) is reported in Consensus Forecasts—namely the average forecast for inflation in country i for the current year, made at time t. There is obviously a similar variable for next year, replacing "c" with "n". Finally, $S_{it}(\pi, c) = S\pi_t^{fi,c}$ is reported in Consensus Forecasts—the standard deviation of the private sector forecasts for the current year for country i at date t. We have 3 observations for each country c at time t: $\overline{\pi}_t^{fi,c}$, $S\pi_t^{fi,c}$,and π_t^{ai}.

We calculate the variance of private sector forecasts = $V_t^{fi,c}$ (across forecasters) = $[S\pi_t^{fi,c}]^2$ as follows:

$$V_t^{fi,c} = E^j[\pi_{jt}^{fi,c} - \overline{\pi}_t^{fi,c}]^2 = E^j[\pi_{jt}^{fi,c}]^2 - [\overline{\pi}_t^{fi,c}]^2$$

We can use this equation to solve for $E^j[\pi_{jt}^{fi,c}]^2$ (since we have all terms on the RHS of the equation):

$$E^j[\pi_{jt}^{fi,c}]^2 = V_t^{fi,c} + [\overline{\pi}_t^{fi,c}]^2$$

We can now calculate the RMSE of the forecast for the "current" year as follows:

$$RMSE_t^{fi,c} = E^j[\pi_{jt}^{fi,c} - \pi_t^{ai}]^2 = V_t^{fi,c} + [\overline{\pi}_t^{fi,c}]^2 + [\pi_t^{ai}]^2 - 2\overline{\pi}_t^{fi,c}\pi_t^{ai}$$

Notice, we have all terms on the RHS of this equation. We can also calculate the RMSE of the forecast for the "next" year as follows:

$$RMSE_t^{fi,n} = V_t^{fi,c} + [\overline{\pi}_t^{fi,n}]^2 + [\pi_{t+1}^{ai}]^2 - 2\overline{\pi}_t^{fi,n}\pi_{t+1}^{ai}$$

where we use actual inflation at "t+1" rather than at "t".

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