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By

Thórarinn G. Pétursson

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Long-term inflation expectations and inflation dynamics

Thórarinn G. Pétursson*
Central Bank of Iceland
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After rising sharply following the Global Financial Crisis, inflation in Iceland has been low and stable in recent years despite a strong cyclical recovery. This not only reflects favourable external conditions but also coincides with a significant decline in long-term inflation expectations in financial markets. It is argued, however, that this market-based measure of inflation expectations actually underestimates the true decline in long-term inflation expectations of price setters. To extract this unobserved wedge between inflation expectations of price setters and financial agents, we estimate a time-varying parameter Phillips curve model for the inflation-targeting period since 2001, adjusting also for an unobserved risk premium in market-based inflation expectations. The empirical results suggest that the expectations wedge was significantly positive until early 2012, after which it starts to gradually decline towards zero. The true decline in long-term inflation expectations of actual price setters is therefore much steeper than is captured by the market-based measure and taking this into account results in a stable and plausible specification of the Phillips curve that can explain key features of the recent inflation developments in Iceland.

JEL codes: E31, E32, E37, E52.

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1. Introduction

Iceland has recently experienced an unusually long period of stable and low inflation. Inflation has averaged 2% since the start of 2014, compared to 5% since the start of the inflation-targeting regime in 2001. The focus of this paper is to analyse the factors behind this recent development. Is it simply "good luck" reflecting the imported deflation stemming from the global "missing inflation" phenomena (cf. IMF, 2016), or does it also reflect an improvement in domestic monetary policy performance manifesting itself in a decline in long-term inflation expectations towards the official inflation target?

To answer this, a forward-looking, open-economy Phillips curve is estimated over the inflation-targeting period. The analysis suggests a structural break in the average relation between inflation on one hand and inflation expectations, cyclical output, and relative import prices on the other hand. The data suggests that this structural break occurred in early 2012 when inflation started easing from over 6% to 1% in early 2015 despite a strong rebound in economic activity and continuous decline in unemployment. The disinflation coincides with a significant decline in long-term market-based inflation expectations (using the 10-year breakeven inflation rate from the bond market) from roughly 5% towards the 2.5% inflation target. It is argued, however that the decline in this market-based measure of long-term inflation expectations (the only one available over the whole sample period) actually underestimates the true decline in long-term inflation expectations of price setters. To estimate this unobserved "expectations wedge", we use a time-varying parameter model, estimating the expectations wedge jointly with an unobserved risk premium in the breakeven inflation rate. We find that the wedge has gradually declined from 1.5 percentage points in 2010 to zero towards the end of the sample period. This suggests a much steeper decline in long-term inflation expectations of actual price setters over the sample period than is captured by the market-based measure of inflation expectations from a peak of almost 8% in mid-2011 to close to the inflation target by the end of 2016.

Taking the gradual decline in the unobserved expectations wedge into account results in a stable and plausible specification of the Phillips curve. The empirical findings suggest that the combination of declining bond market inflation expectations and expectations wedge, together with a large imported deflation, play a key role in explaining the disinflation since 2012 and the continued low inflation despite the strong growth in economic activity. Furthermore, the failure to take the decline in the expectations wedge since 2012 into account goes a long way in explaining the persistent over-prediction of inflation during the disinflation episode. So, to answer the question set out at the beginning: a fair slice of good luck and improved monetary policy credibility have combined to reduce inflation and push long-term inflation expectations towards the inflation target over the last few years.

The remainder of the paper is organised as follows. Section 2 briefly describes the inflation-targeting regime in Iceland and the reforms made following the financial crisis in 2008. Section 3 presents the empirical results for different specifications of the Phillips curve focusing on the recent structural break that appears in the relationship and how that can be attributed to a decline in a wedge between the unobserved inflation expectations of price setters and expectations measured from financial markets. Section 4 discusses the role of this decline in long-term inflation expectations in explaining the recent disinflation period. Section 5 concludes.

2. Inflation targeting and recent monetary policy reforms

Much of Iceland's post-war economic history can be characterised as a period of chronically high and volatile inflation. Various exchange rate and monetary policy regimes were tried, ultimately leading to the introduction of the inflation-targeting regime in March 2001. The early signs seemed promising: inflation reached the 2.5% inflation target in late 2002 and remained close to target until mid-2004. However, macroeconomic imbalances had started to emerge again following the privatisation of the domestic banking system and the ensuing credit boom. Demand pressures mounted yet again and inflation overtook the 4% upper deviation limit in early 2005 and remained above it almost without interruption until the second half of 2010, peaking at almost 20% in January 2009 following a sharp depreciation of the exchange rate once the financial crisis started in full force in late 2008.

The persistent overshooting of the inflation target and the catastrophic financial crisis in 2008 highlighted a number of fundamental weaknesses in the monetary policy framework and shortcoming in the conduct of overall macroeconomic and financial stability policy in Iceland. A number of significant changes to the monetary policy framework were therefore introduced in early 2009 (see Pétursson, 2019 for more details). A single governor replaced a three-member Board (typically headed by a former political leader) and a five-member monetary policy committee with two external members from academia replaced the Board as the monetary policy decision-maker. Policy decisions are reached by a simple majority and are announced at a press conference on the decision day, followed by the publication of minutes two weeks later (which includes information on individual voting). The committee also appears before a parliamentary committee twice a year. This constitutes a significant change from the previous setup. The previous Board only started holding fixed, pre-announced rate setting meetings in 2006, no minutes of policy meetings were ever published, and information on the voting and individual views of Board members was not made available. Public speaking of policy makers explaining the rationale behind policy decisions was relatively rare and there was no fixed structure for parliamentary hearings. The reforms implemented in 2009 therefore significantly enhanced the transparency of monetary policy, with the Dincer and Eichengreen (2014) transparency score showing the biggest gain in monetary policy transparency in Iceland since 2008 among advanced economies. The reforms also saw an extension of the policy toolkit to include various macro-prudential tools and more active use of foreign exchange interventions and capital flow measures to lean against capital flow surges and financial cycles.

The reforms also appear to coincide with less tolerance towards inflation target overshooting than before (see, for example, Central Bank of Iceland, 2017). This became clear in August 2011 when the policy committee decided to raise interest rates after a continuous easing of policy since early 2009. While a sizable slack still remained, the committee was faced with a sharp rise in inflation expectations following a generous wage settlement in the spring of 2011. The committee therefore responded by raising rates and signalled further rate hikes, eventually ending by raising rates by 175 basis points in just over a year until November 2012 when long-term inflation expectations had started to decline again. The negative reaction to the rate-hike cycle from politicians and the population at large was enormous, but it may have strengthened the beliefs of economic agents of the firm intentions of the committee to anchor inflation at the 2.5% inflation target. A further bout of wage inflation came in the spring of 2015, raising long-term inflation expectations significantly again. The committee responded by hiking rates (again to strong popular opposition), perhaps further cementing its inflation-fighting credentials before easing rates back once inflation expectations declined towards the target in late 2016.

These changes to the monetary policy framework, its strategy, communication, and implementation have coincided with significant improvements in inflation performance, with target misses declining significantly towards what is typically observed in other advanced economies (cf. Central Bank of Iceland, 2017). From mid-2015, inflation has fluctuated between 1.5% and 3% for most of the period, and it has remained within the 1-4% deviation range of the inflation target for a longer period than any time before since the start of the inflation-targeting regime. After averaging around 5% over the whole inflation-targeting period – twice the inflation target – trend inflation (measured as the 5-year moving average of inflation) has eased from a peak of 8% in 2009-2012 to below 3% by the end of 2016. There are therefore signs that inflation and inflation expectations have become better aligned with the inflation target and that the credibility of the framework has improved. In the following section we aim to quantify to what extent this improved inflation performance can be attributed to improved anchoring of long-term inflation expectations using a forward-looking, open-economy Phillips curve.

3. Empirical analysis

3.1. A Phillips curve specification of inflation dynamics

The Phillips curve is one of the core building blocks of macroeconomics, linking the nominal and real side of the economy through the relation between inflation and economic activity, as first observed by Phillips (1958). Inflation expectations were given a prominent role in the "expectations-augmented" version of the Phillips curve by Friedman (1968) and Phelps (1968), with the New-Keynesian literature providing the micro-foundations for the relationship through Calvo's (1983) model of staggered price adjustment (see, for example, Gali and Gertler, 1999, and Mavroeidis et al., 2014, for a recent survey of the empirical literature). The New Keynesian Phillips curve (NKPC) can be written as:

$$(1) \qquad \pi_t = \pi_t^e + \kappa x_t$$

where π_t is inflation, π_t^e is inflation expectations, x_t is the output gap, and κ measures the slope of the Phillips curve and is determined by the underlying structural parameters of the model.

To better capture the inherent persistence typically found in inflation rates and to allow us to focus on the role of improved anchoring of long-term inflation expectations to the inflation targeting, we follow Matheson and Stavrev (2013) and assume that inflation expectations are given as the weighted average of past inflation and long-term inflation expectations:

$$(2) \qquad \pi_t^e = \beta \pi_{t-1} + (1-\beta)\pi_t^*$$

where π_t^* is long-term inflation expectations. Furthermore, since inflation is measured with consumer price inflation, we add relative import price inflation to capture the effects of global inflation and exchange rate movements. The Phillips curve then becomes:

$$\pi_t = \beta \pi_{t-1} + (1-\beta)\pi_t^* + \kappa x_t + \lambda \pi_{qt-1} + \epsilon_t \tag{3}$$

where π_{qt} is relative import price inflation (included lagged by one quarter to reduce possible problems related to endogeneity of the regressor). A residual, ϵ_t, is also included that can capture the effects of transitory supply shocks on inflation not reflected in movements in the output gap and relative import prices.

This hybrid specification of the Phillips curve encompasses the simple expectations-augmented ($\beta = 0$) and accelerationist ($\beta = 1$) Phillips curves but, following IMF (2013, 2016), Blanchard et al. (2015), and Blanchard (2016), differs from the typical NKPC specification in that one-period ahead inflation expectations have been replaced with long-term inflation expectations. Variation in long-run inflation expectations can be an important source of inflation dynamics (cf. Kozicki and Tinsley, 2002) and they should be a close proxy for the inflation target that economic agents believe that the authorities are aiming for (see, for example, Del Negro et al., 2015, and IMF, 2016). If the inflation-targeting regime is fully credible, these long-run expectations should correspond to the official inflation target and the coefficient $1-\beta$ capture the degree of "level anchoring" of inflation expectations (cf. Ball and Mazumder, 2011). If, however, the inflation-targeting regime lacks credibility, persistent changes in long-term inflation expectations should reflect the lack of credibility in the inflation target. The specification of the Phillips curve in Eq. (3) therefore seems particularly suitable for an analysis of how possible changes in the credibility of monetary policy in Iceland have affected inflation dynamics in recent years.

Ideally, the measure of long-term inflation expectations used should capture the expectations of actual price setters. However, survey measures of firms' long-term inflation expectations are not available. Neither is data on the long-term inflation expectations of households, which a number of studies suggest are a good proxy for inflation expectations of firms (cf. Kumar et al., 2015, Coibion and Gorodnichenko, 2015, and Coibion et al., 2018). The only measure of long-term inflation expectations available is from the financial market, but these are likely to be an imperfect measure of inflation expectations of price setters. Market participants are, for example, more likely to be paying closer attention to monetary policy actions and signals than the general population and there is evidence that their inflation expectations are more forward-looking than those of households and firms (cf. Coibion et al., 2018). This lack of attention and slow accumulation of information can be rationalised in many ways (see Coibion et al., 2018, for an overview of the main theories). For example, in the sticky-information model of Mankiw and Reis (2002), economic agents slowly accumulate information about macroeconomic variables, and in Carroll's (2003) epidemiological model, "expert opinion" on the inflation outlook spreads slowly through the news media to the general public which absorbs the new information probabilistically. For example, Mankiw et al. (2003) show that following the Volcker disinflation period of the early 1980s, only some economic agents quickly revised their inflation expectations downwards, whereas others retained their pre-Volcker expectations. As the disinflation proceeded, a larger share of the population updated their beliefs and revised their inflation expectations downwards.

Households and firms therefore seem to take a longer time to change their view on the inflation outlook and to be convinced that the inflation regime has in fact changed. This suggests that using inflation expectations from financial market agents could lead to an underestimation of long-term inflation expectations of actual price setters and overstate their decline during disinflation episodes, and that this can be a source of instability in the Phillips curve. To take account of this measurement error, we re-write long-term inflation expectations as:

$$(4) \qquad \pi_t^* = \pi_t^{uc} + \pi_t^m$$

where π_t^m is long-term inflation expectations from the financial market and π_t^{uc} is an unobserved wedge between inflation expectations of price setters and market participants. A decline in this "expectations wedge" would suggest that long-term inflation expectations of price setters have become more aligned with those of financial market participants and a decline in overall long-term inflation expectations π_t^* towards the inflation target would suggest that the inflation-targeting regime has become more credible.

A further complication is that direct measurements of π_t^m are not available (except for the last few years). Therefore, they have to be extracted from the interest rate spread between nominal and inflation-indexed government bonds, which measures the expected inflation over the maturity of the bonds that would make a risk-neutral investor indifferent between holding either of the bonds – typically called the breakeven inflation rate. This measure of inflation expectations can therefore contain a (possibly) time-varying risk premium for inflation and liquidity risks. Thus, if π_t^b is the long-term breakeven inflation rate, we get:

$$(5) \qquad \pi_t^b = \pi_t^m + \rho_t$$

where π_t^m is inflation expectations from the financial market, or the risk-adjusted long-term breakeven rate, and ρ_t is the unobserved risk premium (the sum of risk premia for inflation and liquidity risk). In this case, Eq. (4) becomes:

$$(6) \qquad \pi_t^* = \pi_t^{uc} + \pi_t^b - \rho_t$$

with actual long-term inflation expectations now containing two unobserved components, the expectations wedge, π_t^{uc}, and the bond market risk premium, ρ_t.

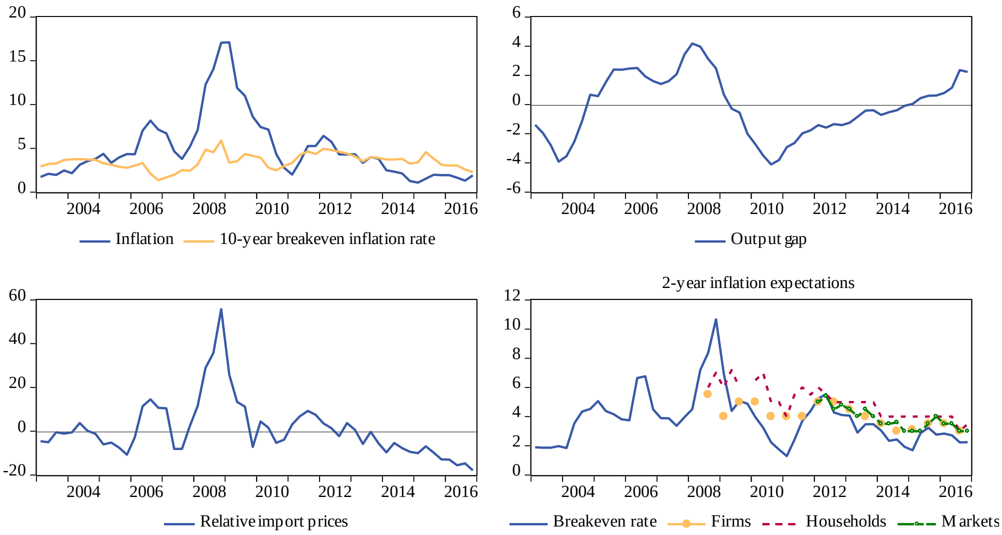
3.2. A first attempt at the Phillips curve

We start by estimating the Phillips curve allowing for a single unobserved component in long-term inflation expectations:

$$(7) \qquad \pi_t = \beta\pi_{t-1} + (1-\beta)(\alpha_t + \pi_t^b) + \kappa x_t + \lambda\pi_{qt-1} + \epsilon_t$$

where $\alpha_t = \pi_t^{uc} - \rho_t$, i.e. the difference between the unobserved inflation expectations wedge and the bond market risk premium. Inflation is measured as the year-on-year change in headline Consumer Price Index (CPI), the output gap is measured as the four-quarter (trailing) moving average of the quarterly difference between actual and potential GDP, and relative import price inflation is measured as the year-on-year change in the ratio of local currency import price deflator to the GDP price deflator. The CPI and the import price and GDP price deflators are obtained from Statistics Iceland, while the data on inflation expectations and the output gap are from the Central Bank of Iceland. We focus on the inflation-targeting period and use quarterly data from 2003 (the start of the breakeven inflation rate data). Figure 1 shows the data while the Appendix reports how robust the results are to alternative data specifications and to alternative modelling and estimation approaches.

Figure 1. The data



Note: The upper panel shows the data for CPI inflation and the 10-year breakeven inflation rate (year-on-year, %) and the output gap (% of potential output). The lower panel shows relative import prices (year-on-year, %) and the 2-year breakeven inflation rate and three different survey measures of 2-year inflation expectations (firms, households, and financial markets) (year-on-year, %).

Source: Central Bank of Iceland, Gallup, Statistics Iceland.

We start by estimating a simple linear version of the Phillips curve, treating α_t as a constant and without imposing the dynamic homogeneity restriction in Eq. (7). The results are reported in the first column of Table 1. The Andrews (1993) sup$F$ test for a structural break at an unknown date strongly suggests a structural break in the Phillips curve relation around 2012. Although it cannot be ruled out that the instability stems from the parameters on the lagged or forward inflation terms, further testing suggests that the instability originates from the α_t term.[1] The second column of Table 1 therefore allows for an unknown number of structural breaks in α_t at unknown dates using the Bai and Perron (2003) sequential estimation approach. The dynamic homogeneity restriction is also imposed (p-value = 0.13). This approach gives a single break point in α_t in 2012Q2, which declines from a statistically significant α_1 = 1.93 percentage points before 2012Q2 to a non-significant α_2 = 0.17 percentage points from 2012Q2. Once taking the downward shift in α_t into account, the Phillips curve specification in Eqs. (7) seems sufficient to capture inflation dynamics in Iceland over the inflation-targeting period, in particular the post-crisis disinflation towards the end of the sample period. Although there is some evidence of serial correlation and

[1] Once a dummy variable is added to the regression to capture the structural break in α_t, further dummy variables on lagged and forward inflation become statistically insignificant.

heteroscedasticity in the residuals in the unrestricted specification of the Phillips curve, this disappears once the homogeneity restriction is imposed and the structural break in α_t is allowed for.

Table 1. Linear specifications of the Phillips curve

	Unrestricted specification with constant α_t		Restricted specification with structural break in α_t	
Parameter	Estimate	Std. error	Estimate	Std. error
α_1	-0.583	0.870	1.933	0.307[a]
α_2	–	–	0.166	0.424
β	0.576	0.071[a]	0.486	0.060[a]
γ	0.702	0.160[a]	–	–
$1-\beta$	–	–	0.514	0.060[a]
κ	0.399	0.063[a]	0.402	0.058[a]
λ	0.096	0.021[a]	0.108	0.017[a]
R^2 (adj.)	0.934		0.942	
σ_ϵ	0.961		0.898	
$\log L$	-74.636		-70.809	
sup$F$ test	p-value	Date		
α	0.001	2011Q4		
β	0.013	2012Q2		
γ	0.001	2011Q4		
κ	0.209	–		
λ	0.367	–		
BG test	0.016		0.144	
White test	0.003		0.070	

Note: The table reports the regression results for two versions of the Phillips curve for the sample period of 2003Q1-2016Q4 (T = 56). a, b, and c denotes point estimates significant at the 1%, 5%, and 10% critical level, respectively. The unrestricted version of the Phillips curve treats α_t in Eq. (7) as a constant without imposing the dynamic homogeneity restriction (with the coefficients on lagged inflation and inflation expectations given as β and γ, respectively). The sup$F$ test is the Andrews (1993) test for a structural break of an unknown date. The p-values reported are obtained using Hansen's (1997) method. The restricted version of the Phillips curve imposes the homogeneity restriction and allows for multiple breakpoints in α_t at unknown dates using the Bai and Perron (2003) sequential $L+1$ vs. L estimation approach. This approach finds a single breakpoint in α_t in 2012Q2, F-value = 10.16 (critical value = 8.58), and no evidence of additional breakpoints, F-value = 9.81 (critical value = 10.13). The table also reports p-values for the Breuch-Godfrey test for first-order serial correlation and the White test for heteroscedasticity.

Source: Author's calculations.

3.3. A Phillips curve with α_t treated as an unobserved random variable

Although Eq. (7) with a downward shift in α_t in early 2012 adequately captures inflation dynamics in Iceland over the sample period, the way that the change in α_t is modelled is not completely satisfactory. The deterministic break in 2012Q2 implicitly implies a sudden shift in expectations formation in early 2012, which economic agents identify as a new regime with a probability of one. This seems implausible when thinking about the general population gradually learning and updating their views on the commitment and credibility of the monetary policy authority as new information on inflation and policy responses becomes available (cf. Backus and Driffill, 1985, and Barro, 1986).

Therefore, instead of treating the movement in α_t as a single deterministic break from 2012, we assume that it is an unobserved continuous random variable, specified as the following random walk process:

$$\alpha_t = \alpha_{t-1} + v_t \tag{8}$$

where v_t is an independent white noise error with variance σ_v^2. We apply the Kalman filter to estimate this unobserved random walk process simultaneously with other parameters of the Phillips curve in Eq. (7). The results are reported in Table 2.

Table 2. Phillips curve with one unobserved component

	Unrestricted TVP model		Restricted TVP model	
Parameter	Estimate	Std. error	Estimate	Std. error
β	0.203	0.105[c]	0.429	0.061[a]
$1-\beta$	0.797	0.105[a]	0.571	0.061[a]
κ	0.441	0.195[b]	0.412	0.073[a]
λ	0.126	0.017[a]	0.114	0.016[a]
α_T	0.582	0.905	0.283	0.657
σ_ϵ	0.487	0.201[b]	0.844	0.077[a]
σ_v	0.750	0.227[a]	0.267	–
$\log L$	-78.038		-80.529	

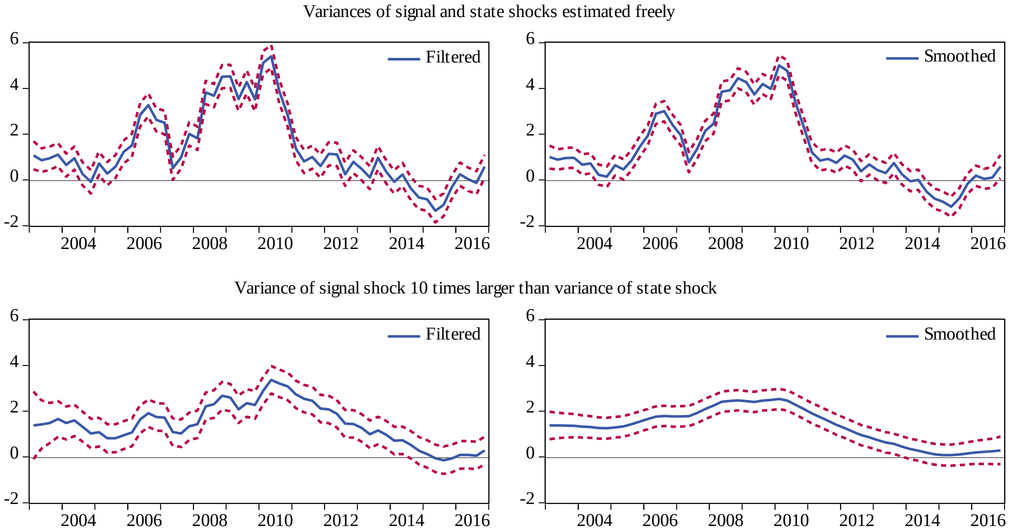
Note: The table reports estimation results for the TVP specification of the Phillips curve in Eqs. (7), with α_t determined by Eq. (8). The sample period of 2003Q1-2016Q4 ($T = 56$). a, b, and c denotes point estimates significant at the 1%, 5%, and 10% critical level, respectively. α_T gives the final state estimate of α_t. The unrestricted TVP model estimates σ_v^2 and σ_ϵ^2 freely, while the restricted TVP model imposes the restriction that $\sigma_v^2 = \sigma_\epsilon^2/10$ on the variance of α_t.

Source: Author's calculations.

The first column gives the parameter estimates of this time-varying parameter (TVP) Phillips curve. It reports the point estimates of the parameters of Eq. (7) that are assumed to be constant and the end-of-sample estimate of α_t in Eq. (8). The time-invariant parameters are very similar to the estimates reported in Table 1, although the weight on lagged inflation is now somewhat lower.

The upper panel of Figure 2 gives the time path of α_t. Two versions are reported: the filtered estimate gives an estimate of α_t using only data up to period t, while the smoothed estimate gives the full sample estimate of α_t. The figures show a clear downward trend in α_t from mid-2010 when it starts declining from a peak of roughly 5 percentage points towards zero; and becoming statistically insignificant from zero by mid-2012.

Figure 2. Time-varying estimate of the unobserved component α_t



Note: Filtered and smoothed Kalman estimates of α_t from Table 2 (in percentage points). The upper panel gives the estimates when σ_ϵ^2 and σ_v^2 are estimated freely while the lower panel gives the estimates when $\sigma_v^2 = \sigma_\epsilon^2/10$. Broken lines show 68% confidence intervals.

Source: Author's calculations.

In addition to the clear downward trend, the figures also exhibit some short-term variation in α_t. In fact, the estimated variation in v_t is larger than in ϵ_t as shown in Table 2. However, it may be more plausible to assume that α_t does not vary greatly from quarter to quarter but is a more slow-moving process. Table 2 therefore also presents a specification of the TVP model assuming that the variance of v_t is one-tenth of the variance of ϵ_t (the restricted variance estimates lie within their unrestricted 95% confidence intervals). The other parameters of the Phillips curve are more or less unchanged, although the size of β increases to just above 0.4 – which is closer to the linear regression estimates reported in Table 1. The estimated time path of α_t after imposing this additional smoothness is given in the lower panel of Figure 2. Similarly to the unrestricted estimate, it rises steadily from 1.5 percentage points at the start of the sample to 2.5 percentage points in mid-2008 (averaging 1.8 percentage points from 2003-2011) before gradually easing to zero (and becoming statistically insignificant from zero in late 2013).

3.4. Adjusting for an unobserved risk premium in breakeven inflation

Up till now, we have ignored the possibility that some of the variation in the unobserved component α_t could be coming from the bond market risk premium, ρ_t. Note that ρ_t enters the definition of α_t with a negative sign. Thus, ρ_t would have to been rising in recent years for it to explain the recent decline in α_t – which may seem implausible given the declining level and volatility of inflation and inflation expectations over the same period. However, as the risk premium also includes a compensation for liquidity risk (the difference between the liquidity risk premia on nominal and inflation-indexed debt) it cannot be excluded a priori that some of the decline in α_t can be explained by a rise in ρ_t instead of a decline in the inflation expectations wedge π_t^{uc}.

To answer this, we need a model for extracting ρ_t from the breakeven inflation rate data in Eq. (5). To do this we use the signal extraction approach suggested by Gurkaynak et al. (2010), which use survey-based inflation expectations data to identify the underlying inflation expectations and risk premium from US bond market data. As we do not have long-term survey data on inflation expectations over most of the period available, we use 2-year median survey responses (the longest horizon survey data available for households and firms) for households (available quarterly since 2008Q3), firms (available semi-annually since 2008Q3), and financial markets (available quarterly since 2012Q1). Together with the 2-year breakeven data (available for the whole period), we can estimate the two unobserved components, π_t^{uc} and ρ_t, jointly with the Phillips curve (the survey data is shown in Figure 1):

(9) $$\pi_t = \beta\pi_{t-1} + (1-\beta)(\pi_t^{uc} + \pi_t^b - \rho_t) + \kappa x_t + \lambda\pi_{qt-1} + \epsilon_t$$

The two other signal equations are now given as:[2]

(10) $$\pi_{2,t}^b = \pi_{2,t}^e + \rho_t$$

[2] This assumes that 2-year inflation expectations of households and firms are equal to those in financial markets, i.e. that there is no wedge in 2-year inflation expectations as it is in 10-year expectations. As shown in Figure 1, the sample averages of these 2-year expectations are similar (in particular for firms and market participants) and a possibly time-varying 2-year wedge is found insignificant when added to the model (see the Appendix for details).

$$\pi_{2,t}^{j} = \pi_{2,t}^{e} + \omega_t^{j} \qquad j = h, f, m \tag{11}$$

with Eq. (10) giving the 2-year breakeven rate, $\pi_{2,t}^{b}$, as the sum of the unobserved 2-year inflation expectations, $\pi_{2,t}^{e}$, and the risk premium, ρ_t, and Eq. (11) specifying the survey-based 2-year inflation expectations of households, firms, and financial markets, $\pi_{2,t}^{j}$ ($j = h, f, m$), as noisy measures of the unobserved 2-year inflation expectations.

There are additional three transition equations, the first two specifying the two expectations terms as random walks, while the final one specifies the risk premium as an AR(1) process:

$$\pi_t^{uc} = \pi_{t-1}^{uc} + u_t \tag{12}$$

$$\pi_{2,t}^{e} = \pi_{2,t-1}^{e} + \psi_t \tag{13}$$

$$\rho_t = \theta\rho_{t-1} + \xi_t \tag{14}$$

The error terms, ϵ_t, ω_t^{j} ($j = h, f, m$), u_t, ψ_t, and ξ_t, are independent white noise errors with variances σ_ϵ^2, σ_j^2 ($j = h, f, m$), σ_u^2, σ_ψ^2, and σ_ξ^2, respectively.

Note that this signal-extraction approach does not give us a direct measure of the 10-year risk premium but only of the unobserved 2-year risk premium. However, these premia are typically found to move closely together although they may differ in levels (cf. Gurkaynak et al., 2010, Chen et al., 2010, and Liu et al., 2015). Our approach should therefore give us a reasonable estimate of the 10-year risk premium, although it may be that we are only able to identify ρ_t (and therefore π_t^{uc}) up to a constant. Note, however, that our estimate of the underlying long-term inflation expectations of price setters, π_t^*, is not affected as that only relies on the estimation of α_t which is not affected by the split between the two unobserved components, π_t^{uc} and ρ_t (with $\pi_t^* = \pi_t^{uc} + \pi_t^{b} - \rho_t = \alpha_t + \pi_t^{b}$).

Table 3. Phillips curve with two unobserved components

Parameter	Estimate	Std. error
β	0.556	0.108[a]
$1-\beta$	0.444	0.108[a]
κ	0.570	0.117[a]
λ	0.117	0.022[a]
θ	0.811	0.108[a]
π_T^{uc}	-0.255	0.875
$\pi_{2,T}^e$	3.000	0.351[a]
ρ_T	-0.616	0.836
σ_ϵ	1.001	0.129[a]
σ_u	0.317	–
σ_ψ	0.282	0.136[b]
σ_ξ	0.700	0.231[a]
σ_h	0.617	0.285[b]
σ_f	0.331	0.208
σ_m	0.000	0.047
$\log L$	-253.941	

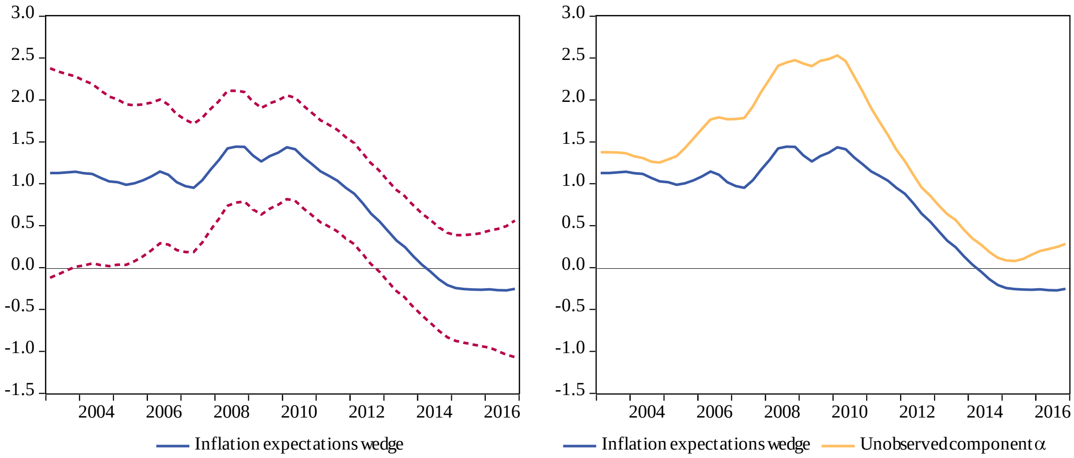
Note: The table reports estimation results for the TVP system in Eqs. (9)-(14). The sample period of 2003Q1-2016Q4 ($T = 56$). a, b, and c denotes point estimates significant at the 1%, 5%, and 10% critical level, respectively. π_T^{uc}, $\pi_{2,T}^e$, and ρ_T give the final state estimates of π_t^{uc}, $\pi_{2,t}^e$, and ρ_t, respectively. The smoothing restriction $\sigma_u^2 = \sigma_\epsilon^2/10$ is imposed on the variance of π_t^{uc}.

Source: Author's calculations.

Table 3 reports the results for Eq. (9)-(14), again imposing additional smoothing on the variance of π_t^{uc}, such that $\sigma_u^2 = \sigma_\epsilon^2/10$. The parameters of the Phillips curve are tightly estimated and are very similar to those previously obtained. The estimate of the autocorrelation coefficient of ρ_t, θ, is found to be close to 0.8, suggesting a significant persistence in the risk premium. The left panel of Figure 3 shows that the smoothed estimate of π_t^{uc} rises from 1 percentage points at the start of the sample to close to 1.5 percentage points in mid-2008, where it remains until mid-2010, before gradually declining to zero (and becoming statistically insignificant from zero in mid-2012).[3] As the right panel of Figure 4 shows, the full-system estimate of π_t^{uc} is similar to the previous estimate of α_t: although the full-system estimate of π_t^{uc} does not increase as much in the run up to the financial crisis as the previous estimate of α_t suggested, the decline since mid-2010 is very similar. The results therefore suggest, as suspected, that the decline in α_t since 2012 is due to the decline in the inflation expectations wedge π_t^{uc} rather than a rise in the risk premium ρ_t.

[3] The wider confidence interval prior to 2008 reflects the lack of 2-year survey data before 2008.

Figure 3. Time-varying estimate of inflation expectations wedge



Note: The left panel gives the smoothed Kalman estimates of the inflation expectations wedge, π_t^{uc}, from Table 3 (in percentage points). Broken lines show the 68% confidence interval. The right panel compares the estimate of π_t^{uc} from Table 3 to the estimate of the unobserved component α_t from Table 2 (smoothed estimate).

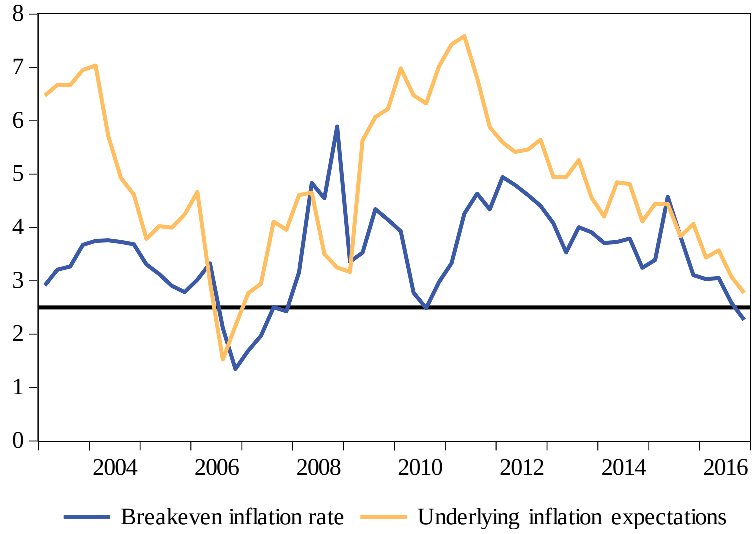
Source: Author's calculations.

4. Long-term inflation expectations and their role in recent inflation developments

4.1. Estimation of long-term inflation expectations

The empirical results reported above suggest that the declining inflation expectations wedge, π_t^{uc}, has been an important feature of the post-crisis inflation dynamics in Iceland. The earlier high value of π_t^{uc} (and inflation expectations in general) captures the legacy of poor inflation control over a long period of time in Iceland and the inherent stickiness in the expectations formation of price setters that led to a much slower adjustment of their expectations than among financial markets participants as inflation performance gradually improved. This implies that long-term inflation expectations of price setters were both higher and more persistent than suggested by bond market expectations. However, price setters appear to have gradually updated their view of the inflation regime as inflation performance has improved, eventually leading to a convergence in their long-term inflation expectations towards those of financial markets – as both have gravitated downwards towards the inflation target.

Figure 4. Breakeven inflation and underlying long-term inflation expectations



Note: Underlying long-term inflation expectations of price setters are given as $\pi_t^* = \pi_t^{uc} + \pi_t^b - \rho_t$, using the Kalman filter estimates of π_t^{uc} and ρ_t from Eqs. (9)-(14) reported in Table 3. The solid horizontal line gives the 2.5% inflation target.

Source: Author's calculations.

This can be seen more clearly in Figure 4. The figure compares the observed 10-year breakeven inflation rate, π_t^b, to the measure of long-term inflation expectations of price setters, $\pi_t^* = \pi_t^{uc} + \pi_t^b - \rho_t$, using the smoothed Kalman filter system estimates of π_t^{uc} and ρ_t from the previous section. As the figure shows, underlying long-term inflation expectations were much higher in the early years of the inflation-targeting regime than is captured by the breakeven inflation rate, but by early 2006 both had reached similar levels. Both increase again in late 2006 but by mid-2008 they diverge again, with π_t^* declining sharply at the onset of the financial crisis, while π_t^b continues to increase, which the empirical analysis in the previous section attributes to a rise in the bond market risk premium rather than a rise in inflation expectations. This reverses from mid-2009 with the breakeven rate markedly underestimating the level of long-term inflation expectations, which peak at almost 8% in mid-2011 following a very generous centralised wage settlement in the spring of 2011 as discussed in Section 2. After the start of the monetary tightening cycle in 2011Q3, long-term inflation expectations start falling again and the difference between underlying inflation expectations and the breakeven rate starts to narrow. By the end of the sample, the breakeven rate had fallen to just below the 2.5% inflation target, while underlying long-term inflation expectations remained slightly above the target at 2.8%. This suggests that the credibility of the inflation target had significantly improved over the period, although there was still some way to go to fully anchor expectations at the target by the end of the sample.

Finally, Table 4 reports the estimate of the Phillips curve in Eq. (3) using π_t^* as a measure of long-term inflation expectations. The parameters are very similar to those previously reported (and the dynamic homogeneity restriction is easily accepted, p-value = 0.58) and, unlike the estimates in Table 1, now the sup$F$ test detects no parameter instability. Using π_t^* as a measure of inflation expectations therefore leads to a stable and plausible specification of the Phillips curve (the second column of Table 4) which can explain key features of the recent inflation developments in Iceland.

Table 4. Linear specifications of the Phillips curve using π_t^*

	Restricted specification with a constant		Restricted specification without a constant	
Parameter	Estimate	Std. error	Estimate	Std. error
α	-0.026	0.265	–	–
β	0.491	0.071[a]	0.492	0.071[a]
$1-\beta$	0.509	0.071[a]	0.508	0.071[a]
κ	0.614	0.078[a]	0.614	0.077[a]
λ	0.134	0.021[a]	0.134	0.021[a]
R^2 (adj.)	0.928		0.930	
σ_ϵ	1.002		0.992	
$\log L$	-77.469		-77.475	
sup$F$ test	p-value	Date		
α	0.930	–		
β	0.373	–		
κ	0.679	–		
λ	0.168	–		
BG test	0.438		0.433	
White test	0.019		0.018	

Note: The table reports the regression results for two versions of the Phillips curve in Eq. (3) for the sample period of 2003Q1-2016Q4 (T = 56) using π_t^* from the TVP system estimate in Table 3 as a measure of inflation expectations. a, b, and c denotes point estimates significant at the 1%, 5%, and 10% critical level, respectively. The sup$F$ test is the Andrews (1993) test for a structural break of an unknown date. The p-values reported are obtained using Hansen's (1997) method. The table also reports p-values for the Breuch-Godfrey test for first-order serial correlation and the White test for heteroscedasticity.

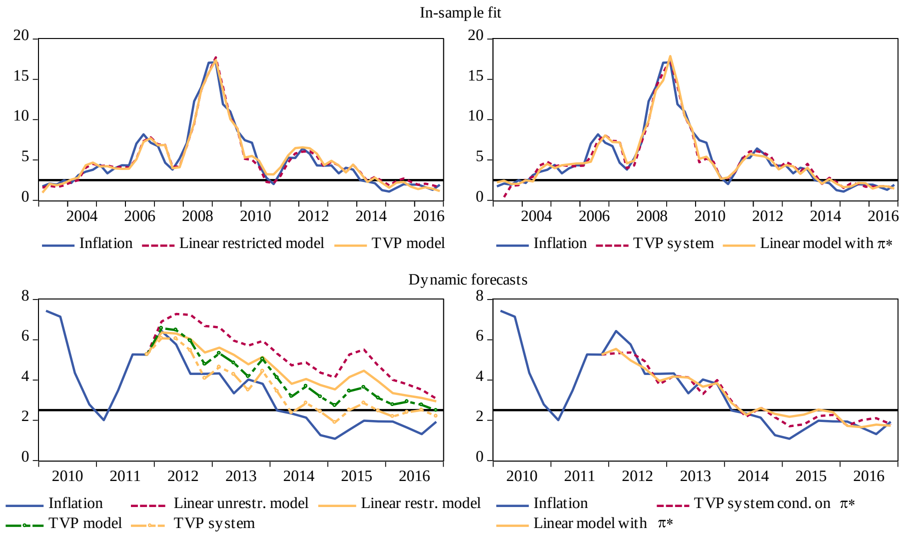
Source: Author's calculations.

4.2. Model fit and forecasting performance

Figure 5 (upper panel) shows the in-sample fit of different specifications of the Phillips curve. All the specifications fit the data well. They capture the gradually rising inflation in the years before the financial crisis and its sharp increase following the currency crisis. They also capture the decline in inflation in 2009 and 2010, although the models predict a faster decline in early 2010 than actually occurred. The temporary increase in inflation in 2011 and early

2012 are also captured; as is the sustained disinflation since early 2012. None of the models fully captures the sharp decline in inflation in late 2014 and in 2015 following the large global oil price shock, however. Thus, it appears that the oil price shock had a larger impact on inflation in Iceland than is captured by the historical link between inflation and relative import prices.

Figure 5. Fit and forecasting ability of different Phillips curve specifications



Note: The upper panel shows actual and fitted inflation, while the lower panel shows dynamic forecasts for the period 2012Q1-2016Q4 for the models re-estimated up to 2011Q4 (all in %). Results are shown for the linear unrestricted model (Table 1, first column), the linear restricted model (Table 1, second column), the TVP model (Table 2, second column), and the TVP system (Table 3). Dynamic forecasts are also shown for the TVP system and the linear model from Table 4 using π_t^* as an observed measure of inflation expectations. The solid horizontal line gives the 2.5% inflation target.

Source: Author's calculations.

The measure of fit reported in the upper panel of Figure 5 essentially represents in-sample one-quarter-ahead forecasts of inflation. However, a more challenging and informative test of these models are dynamic forecasts, i.e. forecasts that use dynamic simulations beyond the estimation period treating past inflation as endogenous (but using observed values of the forcing variables). For this exercise, the models are therefore re-estimated up to 2011Q4 and dynamic forecasts are generated for the five-year period 2012Q1-2016Q4. The results are shown in the lower left panel of Figure 5. As the models are only estimated until 2011Q4, the forecasting improvement of the restricted linear specification over the unrestricted specification stems only from the dynamic homogeneity restriction. The TVP specification from Table 2 improves the forecasting ability of the Phillips curve further still, but the best performing forecasting model is the TVP system from Table 3, although some positive bias remains in the period after 2014.

To what extent does this over-prediction reflect the failure of the Phillips curve specifications to take into account the improved anchoring of inflation expectations as captured by the decline in the inflation wedge π_t^{uc} reported above? To answer this, the dynamic forecasting exercise for the linear specification in Table 4 and the TVP system from Table 3 is repeated but treating π_t^* as an observed variable. The lower right panel of Figure 5 shows that although some of the over-prediction after the oil price shock in late 2014 and early 2015 remains, the positive bias is reduced significantly. The improvement in the forecasting ability of the models once the decline in π_t^{uc} is taken into account can also be seen in the decline in average bias and RMSE in Table 5. This counterfactual exercise therefore suggests that an important part of the apparent over-prediction of inflation in Iceland during the disinflation episode can be explained by the failure to take the gradual improvement in monetary policy credibility since 2012 fully into account. In the next section, we attempt to take a closer look at this issue and quantify to what extent the decline in long-term inflation expectations can explain the disinflation in Iceland since 2012.

Table 5. Forecast errors for different Phillips curve specifications

Different Phillips curve specifications	Bias	RMSE
Unrestricted linear model	2.384	2.503
Restricted linear model	1.590	1.732
TVP model	1.164	1.282
TVP system	0.539	0.736
TVP system conditional on full-sample estimate of π_t^*	0.134	0.546
Linear model with π_t^* as inflation expectations	0.167	0.568

Note: The table reports the average bias and RMSE for dynamic forecasts for the period 2012Q1-2016Q4 for the models reported in Tables 1-3 re-estimated up to 2011Q4 (in percentage points). The table also reports a dynamic simulation for the linear Phillips curve from Table 4 using π_t^* as inflation expectations and the TVP system treating π_t^* as an observed variable.

Source: Author's calculations.

4.3. How large a role did the decline in long-term inflation expectations play during the disinflation episode?

Recent years have seen exceptionally low inflation in all advanced economies, with inflation even starting to decline further from 2011 despite a sustained and synchronised recovery of global demand and historically low unemployment rates in many countries. In fact, inflation rates were typically lower five years after the Global Financial Crisis (GFC) than immediately after the crisis. The fact that this phenomena is so widespread suggests that this "missing inflation" is to an important extent driven by common factors, such as the persistent output slack in most advanced economies, large positive supply shocks (e.g. the large decline in oil, commodity and telecommunication prices), and a sizable overcapacity in the manufacturing sector in a number of large exporting countries (IMF, 2016). IMF (2016) also suggests that available estimates of inflation expectations may underestimate their true decline and that the output slack may actually be larger in many advanced economies than is currently estimated using standard methods (see also Constancio, 2015). This raises the question to what extent the disinflation in Iceland from 2012 and the low recent inflation, despite a strong cyclical recovery, is due to these global factors (and therefore "good luck") and to what extent it is driven by improved credibility of monetary policy as suggested by the decline in long-term inflation expectations.

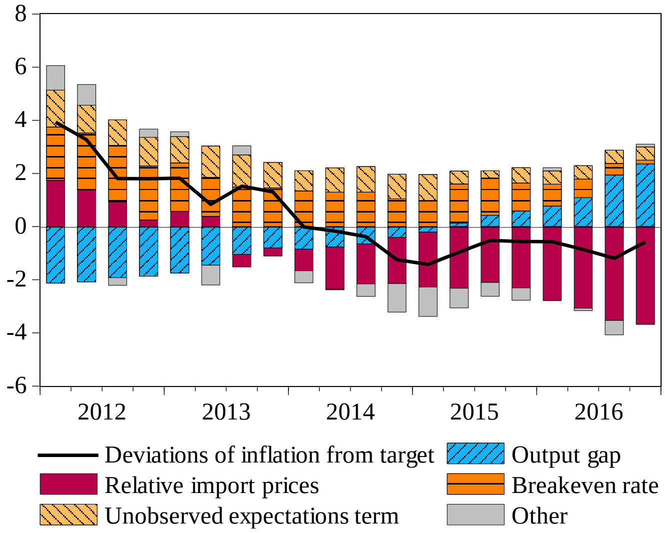
To answer this, we attempt to decompose the inflation dynamics to quantify the contribution of individual factors (inflation expectations, the output gap, and relative import prices) to the development of inflation in recent years. For this exercise, we use the full-sample estimate of the TVP system specification of the Phillips curve from Table 3 to generate dynamic simulations from 2011Q1. The deviations of the simulated inflation paths from the 2.5% inflation target are then decomposed into contributions from each explanatory variable which are constructed from counterfactual simulations where each factor is set to zero (the output gap, relative import prices, and the inflation expectations wedge, π_t^{uc}) or equal to the inflation target (the risk-adjusted breakeven inflation rate, $\pi_t^m = \pi_t^b - \rho_t$).[4] The sum of the contribution from the inflation expectations wedge and the risk-adjusted breakeven rate therefore give the total contribution of long-term inflation expectations, $\pi_t^* = \pi_t^{uc} + \pi_t^b - \rho_t$, to the decline in inflation since 2012.

Figure 6 presents the results. As the figure shows, inflation was significantly above target in the early part of the period, which can mainly be explained by long-term inflation expectations well above target and the lagged effects of past exchange rate depreciation, which overwhelm the negative contribution from the sizeable slack in the economy following the financial crisis. The contribution from the poorly anchored inflation expectations is large: high long-term bond market expectations, as measured by the risk-adjusted breakeven rate, add 2.4 percentage points on average to inflation in 2012, with a further 0.8 percentage point contribution from π_t^{uc}. In total, the deviations of long-term inflation expectations from target

[4] See Yellen (2015) and IMF (2016) for similar exercises. Note that since the simulations are dynamic, each factor also affects inflation through the lagged inflation term. Note, however, that the counterfactual simulations hold other factors constant, thus ignoring the possible effects of the factors on each other.

therefore add more than 3 percentage points to inflation in 2012. The effect starts to wane as inflation expectations decline towards target. The contribution of the expectations wedge disappears by mid-2014 and the overall contribution of inflation expectations gradually declines to 0.6 percentage points by the end of the sample period. Although above-target expectations continue to add to inflation at the end of the sample period, the sizable decline in expectations since 2012 appears to have played a key role in the gradual decline of inflation to target.

Figure 6. Contributions to deviations of inflation from target



Note: The columns give the contribution of each explanatory variable in the Phillips curve to the deviations of inflation from target in 2012Q1-2016Q4 for the TVP system in Table 3 (in percentage points). The contributions are obtained by comparing a dynamic simulation of inflation starting in 2011Q1 to a counterfactual simulation setting the value of the explanatory variable to zero (the output gap, relative import prices, and the inflation expectations wedge, π_t^{uc}) or equal to the inflation target (the risk-adjusted breakeven inflation rate, $\pi_t^m = \pi_t^b - \rho_t$).

Source: Author's calculations.

As Figure 6 shows, import prices start to add further downward pressures on inflation from mid-2013, adding to the impact from the slack in the economy and the declining positive contribution from above-target inflation expectations. These downward pressures from import prices gather strength from 2014, gradually pushing inflation below target. Inflation falls further below target as the combined effect of falling global commodity prices and currency appreciation gathers strength. Partially offsetting this is the gradual disappearance of spare capacity in early 2015 and the consequent emergence of a positive output gap.

Finally, the model residuals ("other" in Figure 6) seem to play a relatively minor role in explaining the deviation of inflation from target in recent years. The exception is in late

2014 and in 2015 when the models fail to explain the full extent of the decline in inflation, reflecting the large oil price shock discussed above. Other than that, the TVP specification of the Phillips curve seems to perform well in explaining the downward trend in inflation and the persistent undershooting of the target since 2014. In that respect, there does not appear to be any puzzle concerning the developments of inflation in the post-GFC period in Iceland.

5. Conclusions

Inflation in Iceland has declined substantially from its high level immediately after the Global Financial Crisis and has remained low and stable for a longer period than seen in decades, despite a strong recovery in domestic activity and a sharp fall in unemployment in recent years. The first disinflation phase, which saw inflation decline from almost 20% in early 2009 to the 2.5% inflation target in early 2011, was mainly driven by the deep recession following the financial crisis, which overwhelmed the lingering effects of the currency collapse in 2008. However, inflation started to rise again in 2011 and reached more than 6% in early 2012. It averaged around 4% from mid-2012 until end-2013 before easing to the inflation target in early 2014. Since then it has mostly remained between 1.5-3%, despite output growth averaging around 4% since 2011, rising to more than 7% in 2016, and unemployment falling below 3% in early 2015.

What explains this? One obvious factor is the low global inflation in recent years that, together with the appreciation of the domestic currency, resulted in a large fall in import prices. However, the analysis presented in this paper suggests that there is more to this than just good luck. A key factor has been the large decline in long-term inflation expectations, which had remained stubbornly well above the inflation target, reflecting the legacy of poor inflation control in Iceland for most of the inflation-targeting period. This improvement in the credibility of the inflation target can be seen in the gradual decline in long-term market-based inflation expectations from around 5% in 2011 to the target in late 2016. This follows widespread reforms of the inflation-targeting framework and its communication and coincides with the start of the first tightening cycle after the financial crisis.

Our analysis suggests, however, that the fall in market-based inflation expectations underestimates the true decline in long-term inflation expectations of actual price setters and a failure to take this into account can explain the apparent structural break in the average relation between inflation and its key drivers in standard Phillips curves occurring in early 2012. We use a time-varying parameter Phillips curve specification to estimate this wedge between unobserved inflation expectations of price setters and market participants, taking into account the unobserved risk premium embedded in breakeven inflation rates. The results suggest a much sharper decline in long-term inflation expectations of actual price setters over the sample period than is captured by the market-based measure of inflation expectations. Further analysis suggests that the failure to take this into account goes a long way in explaining the persistent over-prediction of inflation in Iceland during the disinflation episode. Furthermore, the overall decline in long-term inflation expectations, together with the large imported deflation, play a key role in explaining the post-2012 disinflation. The paper therefore highlights the important role of monetary policy credibility and anchoring of long-term inflation expectations for a successful disinflation and for maintaining low and stable inflation over a sustained period.

Appendix: Robustness analysis

This Appendix explores how robust the results reported in the main text are to various changes in the data specifications and to alternative modelling and estimation approaches. Further detail on the results are available upon request.

1. Data specifications

Output gap

The output gap measure used here is the official Central Bank of Iceland estimate, which uses a production function approach taking into account possible cyclical effects on aggregate supply coming from the cross-border movement of labour and fixed capital over the business cycle, and the impact of the financial crisis in 2008 on resource allocation.

Output gaps are notoriously difficult to estimate and a further complication in Iceland is that production accounts are not available for Iceland at the quarterly frequency. The output gap estimation is therefore based on the expenditure accounts which introduces a strong seasonality and irregular high-frequency noise into the quarterly data. The traditional approach in modelling inflation-output dynamics in Iceland has therefore been to use a four-quarter moving average of quarterly output gap data. This is followed here. Using the raw quarterly data (seasonally adjusted) introduces greater volatility into the output gap measure but all the main results continue to hold. The parameter estimate of the slope of the Phillips curve declines but remains statistically significant, while other parameter estimates remain virtually identical to those reported in the main text. The identification of the signal-to-noise ratio in the unrestricted TVP model in Table 2 deteriorates, however, but the results become almost identical to those reported in the main text once the smoothing restriction is imposed. The same applies when deviations of output from its Hodrick-Prescott trend is used as a measure of the output gap.

We also tried adding the change in the output gap to the Phillips curve to capture potential speed-limit effects but these effects were found to be statistically insignificant. So too were possible additional effects of a global output gap (using the deviations of trade-weighted foreign output from its Hodrick-Prescott trend) to capture the increasing relevance of globalisation on domestic price setting (cf. Borio and Fillardo, 2007).

Relative import prices

Following IMF (2016), we use the ratio of the GDP price deflator and import prices from the national accounts as a measure of relative import prices. An alternative measure is to use the ratio of consumer prices over import prices. Using that measure gave practically identical results to those reported in the main text.

2. Alternative model specifications and estimation approaches

Allowing for an additional wedge in 2-year survey-based inflation expectations

The TVP system was also estimated allowing for an additional 2-year time-varying expectations wedge in the survey-based inflation expectations of households and firms in Eq. (11), i.e. assuming that $\pi_{2,t}^{j} = \pi_{2,t}^{uc} + \pi_{2,t}^{e} + \omega_t^j$ and adding another transition equation for the 2-year expectations wedge, $\pi_{2,t}^{uc} = \pi_{2,t-1}^{uc} + u_{2,t}$. The resulting 2-year wedge shows very limited time variation and is insignificant from zero over the whole sample period. Furthermore, the estimates of π_t^{uc} and ρ_t are virtually identical to those reported in the main text.

Adding short-term oil and commodity price shocks

Although oil prices and non-oil commodity prices are already included in the relative import price variable (with most of oil and non-oil commodity goods imported), the forecast errors after the large oil price shock in late 2014 could suggest that there are additional effects of oil prices that are not fully captured by relative import prices. We therefore tried including short-term oil and non-oil commodity price shocks directly as additional explanatory variables (using quarter-on-quarter changes in these variables). Neither were found to be statistically significant.

Allowing for time variation in all parameters

Although the regression analysis allowing for a structural break in the Phillips curve in Table 1 suggests that changes in α_t are the source of the instability and that other parameters of the Phillips curve can be treated as constant, one could also estimate a TVP model where all the parameters are allowed to vary over time. The marked decline in the inflation wedge continues to come out clearly but no clear pattern or trend can be found in the other parameters, although there is some short-term variation, especially around the financial crisis. However, given their standard deviations, they can be treated as constant throughout the sample period.

Estimating the Phillips curve with instrumental variables

As both inflation expectations and the output gap enter the Phillips curve contemporaneously, the standard parameter estimates could be biased if they are correlated with the equation's residual (cf. Mavroeidis et al., 2014). To test how robust the parameter estimates are to a potential endogeneity problem, the Phillips curve was therefore re-estimated using instrumental variables (using lagged values of the explanatory variables and a shift dummy variable that equals one from 2012Q2 and zero before that as instruments). The resulting parameter estimates continue to be statistically significant from zero and are virtually identical to those reported in the main text. The evidence of a structural break in 2012 also continues to come through. Unlike in Mavroeidis et al. (2014), the instruments appear to strongly identify the parameters of the Phillips curve (using the Cragg-Donald test) – presumably reflecting the fact that inflation in Iceland is both more volatile and predictable than in the US dataset they analyse. Finally, the Wu-Hausmann test suggests that the regressors can safely be treated as exogenous.

Estimating the structural break with a Markov switching model

In an earlier version of the paper (Pétursson, 2018), the unobserved component α_t was also estimated as an unobserved state-dependent variable that can switch between two regimes using a Markov switching approach. The results are very similar to those reported in the main text. In particular, the results suggest a gradual increase in the probability being in the low-inflation regime gradually from practically zero to one from early 2012, with the speed of adjustment very similar to the TVP estimates reported in the main text.

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