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Contents

Opening remarks

Keynote speech

Contributed papers

Programme

18 August 2019, Sunday

19:00	Welcome dinner

19 August 2019, Monday

08:45–09:00	Welcome coffee
09:00–09:15	Opening remarks by BSP Governor Benjamin E Diokno
09:15–09:30	Photo session
Session I	Chair: Dennis Lapid (BSP)
09:30–10:50	Paper 1: What drives inflation in advanced and emerging market economies? Güneş Kamber (IMF), Madhusudan Mohanty (BIS) and **James Morley** (University of Sydney) Discussant: Hans Genberg (Asia School of Business)
10:50–11:10	Coffee break
11:10–12:30	Paper 2: Robots and labour: implications for inflation dynamics **Ippei Fujiwara** (Keio University) and Feng Zhu (Ant Financial) Discussant: Yong Sung Chang (Seoul National University)
12:30–13:50	Lunch
13:50–14:50	Keynote address Chair: Benoît Mojon (BIS) Speaker: Ricardo Reis (London School of Economics and Political Science) Title: The anchoring of long-run inflation expectations today
14:50–15:20	Coffee break

Session II	Chair: Jason Wu (Hong Kong Monetary Authority)
15:20–16:40	Paper 3: The pass-through from short-horizon to long-horizon inflation expectations **James Yetman** (BIS) Discussant: Masazumi Hattori (Nihon University)
16:40–18:00	Paper 4: Can an ageing workforce explain low inflation? **Benoît Mojon** (BIS) and Xavier Ragot (OFCE and Sciences-Po) Discussant: Kenichi Sakura (Bank of Japan)
18:30	Conference dinner

20 August 2019, Tuesday

08:15–08:30	Welcome coffee
Session III	Chair: Ilhyock Shim (BIS)
08:30–09:50	Paper 5: Strategic complementarity and asymmetric price setting among firms **Maiko Koga** (Bank of Japan), Koichi Yoshino (Bank of Japan) and Tomoya Sakata (Bank of Japan) Discussant: Martin Berka (Massey University)
09:50–11:10	Paper 6: Impact of relative price changes and asymmetric adjustments on aggregate inflation: evidence from the Philippines **Joselito Basilio** (BSP) and **Faith Cacnio** (BSP) Discussant: Renée Fry-McKibbin (Australian National University)
11:10–11:30	Coffee break
11:30–12:50	Policy panel on "Central bank policy under changing inflation dynamics: challenges for Asia-Pacific" Chair: Francisco Dakila Jr (BSP) Panellists: Christian Hawkesby (Reserve Bank of New Zealand) Hyungsik Kim (Bank of Korea) Benoît Mojon (BIS)
12:50–13:00	Closing remarks by Benoît Mojon (BIS)
13:00–14:30	Farewell lunch

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Opening remarks

Benjamin E Diokno[1]

A very good morning to all and welcome to this year's Annual Research Conference.

Bangko Sentral ng Pilipinas (BSP), which is currently celebrating the 70th year of central banking in the Philippines, is pleased to co-host this conference with the Bank for International Settlements.

The theme for this year is "Inflation dynamics in Asia and the Pacific". For central bankers, this provides an opportunity to discuss and think about issues that have direct implications for the core mandate of central banks. For academics and research economists, this is a policy-relevant research area that may be openly shared and discussed.

The Great Financial Crisis (GFC) that started in 2008 challenged conventional monetary policy, pushed many central banks outside of their comfort zones, and dramatically altered the landscape of monetary policy. While it was a challenging time for policymakers, it provided fertile ground for macroeconomic research.

A particular topic which received much attention among macroeconomists and central bankers both after and during the GFC was inflation dynamics. An advanced Google search for the phrase "inflation dynamics" for the pre-crisis period of 1997–2007 yielded 4,490 articles. For the post-GFC period of 2009–19, the search generated 13,000 articles – an increase of almost 190%.[2]

Of course, we are all well aware of the challenges faced by central banks in advanced economies (AEs) after the GFC; that despite rock-bottom interest rates and massive liquidity injections, inflation in their economies remained muted. Why this was so was the subject of many studies.

But perhaps one of the reasons for the rapidly rising number of studies on inflation dynamics is the remarkable decline in global inflation which began even before the GFC. For the AEs, this appears to have started in the 1990s, and for the emerging market economies (EMEs), the trend became apparent in the 2000s.

On a regional basis, Asian countries show lower inflation trends among EMEs.[3] Studies reveal that a confluence of both structural and policy-related factors have contributed to this trend in the region.

These include the marked increase in and changing patterns of international trade, especially with the accession of China to the World Trade Organization; technological innovation; trade and financial market integration; demographic transition; structural reforms in labour and product markets that led to greater competition and lesser price rigidities; and more resilient policy frameworks in some

[1] Governor, Bangko Sentral ng Pilipinas.

[2] The search excluded patents and citations, and was conducted on 24 July 2019.

[3] Based on the IMF *World Economic Outlook* April 2019 database.

emerging market and developing economies – particularly the successful adoption of inflation targeting as well as more effective and transparent exchange rates, and monetary and fiscal policy frameworks.[4]

For central bankers, understanding inflation dynamics in relation to other economic variables is crucial. Most, if not all, central bankers will agree with me that controlling inflation is a demanding task. Yes, the central bank can influence inflation by tightening and easing monetary policy. However, the transmission process involves long and variable lags and a degree of uncertainty amid the continuously evolving structure of an economy. Add to this the increasing influence of global and regional factors on domestic inflation. Given these challenges, central banks need to adopt a robust communication strategy to effectively anchor inflation expectations.

As a case in point, the Philippines' headline inflation decelerated from an average of 15.1% in the 1980s to 9.7% in the 1990s. This slowed further to 3.8% from 2002 – when the country adopted the inflation targeting framework – to 2018.

The implementation of the inflation targeting framework in the last 16 years has enabled BSP to manage inflation and to keep it within manageable bounds. This has provided support to greater economic activity, despite international and domestic shocks.

Moreover, the improved transparency, greater accountability and more effective communication resulting from the implementation of inflation targeting have enhanced BSP's credibility. This has led to the continued anchoring of market inflation expectations to monetary policy decisions.

However, in the first few years after the GFC, BSP experienced pressure in its monetary operations as short-term market interest rates diverged from the BSP policy rate. This was mainly due to strong capital inflows arising from the highly accommodative monetary policies and extraordinary liquidity support of central banks in AEs, the steady stream of remittances from overseas Filipinos, and revenues from business process outsourcing.

To address this divergence, BSP launched the interest rate corridor system in June 2016 to enhance the transmission of monetary policy by guiding money market interest rates towards its policy rate.

At the time of writing, this objective has already been achieved as the market rate is now moving in line with BSP's policy rate. In the long term, recalibrations and additions to its monetary policy tools will strengthen BSP's influence on the demand for and supply of money. I am pleased to note that the recently enacted amendments to the BSP charter have restored our authority to issue our own debt securities (even during normal times), which further enhances our monetary policy toolkit.

Further, we will continue to pursue the reduction of reserve requirements from the current 16% to single-digit levels by 2023 to promote a more efficient financial system. This is part of BSP's broad financial sector reform agenda.

We expect that all these actions will aid in the further development of Philippine capital markets by fostering money market transactions and active liquidity management by Philippine banks.

4 See J Ha, A Kose and F Ohnsorge, "Inflation in emerging and developing economies: evolution, drivers and policies", *World Bank Publications*, March 2019.

Indeed, BSP's monetary policy framework has been an effective tool in mitigating the risks emanating from home and abroad. However, we recognise that as the domestic and world economy evolves, all our tools need to be kept constantly appraised. We continually take measures to further enhance the implementation of our inflation targeting framework by improving our array of models, data and software, and the capabilities of our technical staff and modellers.

We likewise ensure prudent exercise of judgment by using all relevant data in our policy decisions. Moreover, we recognise the value of being able to communicate the basis of BSP's policy decisions in a manner that is easy for the public to understand.

All these concepts, and more, will be the subject of this conference's discussions in the next two days. We can rely on an eminent group of economists, who have joined us to present their work and examine issues concerning inflation dynamics and monetary policy.

As central bankers and policymakers, we appreciate this effort to continuously expand our understanding of inflation dynamics and its implications for monetary policy. This is especially so amid the current global economic environment, which continues to shift with the ongoing policy uncertainties and structural changes.

I wish you all a very fruitful conference and a pleasant stay in Manila.

The anchoring of long-run inflation expectations today

Ricardo Reis[1]

Monetary policymakers today benefit from having earned a capital of prestige in the eyes of the public. This capital, and the political clout that comes with it, has allowed them to stay independent even in polarised political times. In part, this capital was earned during the response to the Great Financial Crisis as a new Great Depression was avoided. An even greater part of this public respect has come from policymakers' success at taming inflation. In most OECD countries, inflation is not at the top of the list of concerns that citizens express in surveys. This was not the case in the 1980s and 1990s. But by the end of the 20th century, the variability of long-run inflation had significantly and persistently declined across most advanced countries.

A simple way to illustrate this is to estimate a Beveridge-Nelson model where annual inflation is the sum of a random-walk permanent component and a white-noise transitory component. Doing so for the United States and the euro area shows that the permanent component has been steady since 2000 near the 2% inflation target, and that there has been a clear decline in the estimated variance of this permanent component. In terms of reduced-form statistics, this accounts for the visible fall in the serial correlation of inflation, as well as for the fall in the variance of inflation itself. The permanent component, which one might better call long-run inflation, has become tightly anchored around 2%. This accomplishment is rightly hailed as the proof of success of having an independent central bank with an inflation target.

About 20 years after this change occurred, how do things look today? That is, what is the current inflation anchor? Inflation itself in 2019 may be somewhat above or below target in different countries, but to what long-run level is it converging? This is the topic of this talk. The key emphasis is on long-run inflation, measuring it and controlling it, not on the fluctuations around this anchor. I will repeat the words "long-run" as many times as I can to ensure that this focus is not forgotten.

1. What is the long-run goal of the central bank?

Uncontroversially, I will contend that the long-run goal of the central bank is to control inflation around a stated target. Perhaps a little more controversially, I will argue that there is a good case for this to be the sole goal. Certainly more controversially, I suggest that perhaps the most adequate way to express this target is in terms of a price level target.

[1] Arthur Williams Phillips Professor of Economics, Department of Economics, London School of Economics and Political Science.

The first argument for inflation being the long-run goal of the central bank is simply that this is what is in its legal mandate in most advanced economies. One might wish it was not so, but for a central bank to be legitimately independent it must stick to what the State has instructed it to do.

A second argument is that the central bank can control inflation in the long run. After all, inflation is a monetary phenomenon, so the monetary authority should be the one controlling it.

More interesting is the question of whether the central bank should combine its long-run target for inflation with a long-run target for some measure of real activity, like the unemployment rate or the growth rate of real output. The argument for a long-run dual mandate is that it is hard to reject the null hypothesis that the Phillips curve is not vertical in the data. In fact, when inflation exceeds 30% (or thereabouts), there seems to be a negative relation between the change in the price level and the growth rate of output. There are several theoretical arguments for why a long-run trade-off would exist, including changes in the bargaining power of workers and firms when inflation rises, hysteresis through skills and effort on labour force participation, and the anchoring of expectations on past experience. If such a trade-off exists, why focus solely on inflation?

First, for the higher inflation experiences – well into the two digits – the evidence suggests that lowering inflation increases real outcomes. But then, there is no conflict between a target for inflation and a target for real growth. A redundant dual mandate is better stated as a single mandate.

Second, when inflation is lower, and in the one-digit range, it is also very hard to reject the null hypothesis that the long-run Phillips curve is vertical. The problem is simply that one needs data for a long enough period, and one during which the monetary policy regime is approximately unchanged, in order to estimate the slope of this long-run Phillips curve. We simply do not have this, so any estimates have wide confidence intervals around them. It is hard to move the prior that the central bank can do nothing about real outcomes in the long run.

Third, perhaps this is the right prior to have. Milton Friedman's 1969 presidential address convincingly argued that it is so. Ultimately, the real effects of monetary policy come from fooling agents into confusing absolute for relative price changes. Systematic long-run policy should not rely on constantly fooling people.

Fourth and finally, as a policy stance, I am yet to see a single country successfully exploit a long-run trade-off between inflation and real activity through monetary policy. The principle that if one does not know what the effects of a policy are, one had better not use it (a "do no harm" principle) supports sticking to the prior that monetary policy cannot affect real variables in the long run. As such, output should not be part of the long-run goals of the central bank.

Another goal to consider is financial stability. Yet again, the difficulty is to see where the trade-off is in the long run. High inflation tends to be associated with financial instability. Having a goal of low and stable inflation in the long run is probably what maximises financial stability in the first place.

Unexpected inflation leads to redistribution among different actors that sign nominal contracts and have nominal portfolio positions. Contracts that specify nominal payoffs have their real outcomes depend on what the future price level will be. Therefore, from the perspective of financial stability, it is important to lower the variance of the price level at long-run horizons.

The common practice of inflation targeting, whereby bygones are bygones, imparts a unit root to the price level, which will make this variance grow with the horizon. This suggests that, in order to deliver financial stability, a better practice would be to instead have a target for the price level, which corrects past positive (negative) deviations of inflation from target with negative (positive) deviations in the future, in order to return to the target path for the price level.

This is not such a radical proposal, since the ECB's policy of targeting inflation on average over the medium run, often stated as a 2% target over a five-year horizon, gets very close to what a price level target would be. In the United States today, price level targeting is making a comeback under the new name of average inflation targeting.

A further argument for a price level target is that it provides a commitment to not let below-target inflation persist. It may help central banks regain the credibility they may have lost by repeatedly undershooting their target.

To conclude, the answer to the question I posed is: the long-run goal of the central bank should be to have inflation over the long run stay anchored at its announced target. Stating this in terms of a price level target over, say, a five- or 10-year horizon lowers the forecasting mistakes that agents may make about nominal variables and helps to deliver a better functioning real and financial economy.

2. The centrality of expectations in the long run

To think about the long run, the classical dichotomy is a reasonable starting point. As David Hume famously wrote, a doubling of the amount of money in a person's pockets that comes with a doubling of every price in the economy, should not lead to any change in any actions by any private economic agent. One compelling way to explain this point is that I can count the money in my wallet in dollars, or I can do so in cents, and so can a shopkeeper quote me prices in dollars or cents. It makes no difference which one it is.

In the long run, under the classical dichotomy, inflation is tightly linked to expected inflation. Through multiple channels, expecting higher long-run inflation leads to higher long-run inflation.

If households expect higher inflation, they will hold less currency. This decline in the demand for currency will, ceteris paribus, lead to a fall in the real value of the currency and so to inflation.

If firms expect higher inflation, they will set a rising path for the prices of their goods. As all do so, this results in higher inflation.

If workers expect higher inflation, they will demand higher wages. This raises the nominal marginal costs of firms and leads to higher prices set, and thus inflation.

If investors in financial markets expect higher inflation, they will want to hold fewer reserves at the central banks for a given nominal return on these reserves set by the central bank. This will lower the real value of these reserves, which is the same as higher inflation.

Controlling inflation in the long run then requires controlling expected long-run inflation. Expectations are central to achieving the long-run goal since if they deviate from this goal, bringing inflation and expectations back on target is quite costly.

In the case where expected inflation is too high, experience shows that reducing actual inflation can only happen through a deep recession. In part, this is because adaptive expectations imply that private agents keep on expecting high inflation even after it has started declining. The unexpected decline then becomes associated with some firms setting too-high prices, some workers asking for too-high wages, and some households choosing too-high savings, all inducing a contraction in real activity. Another reason is that once agents expect high inflation, indexation clauses become the norm in many contracts, especially in the labour market. Lowering these built-in inflation clauses requires renegotiations that often are only triggered by rising unemployment. Finally, lowering inflation when expectations are high often requires a fiscal reform that provides fiscal backing to the central bank in its efforts to engage in contractionary monetary policy. This reform is tied to fiscal austerity, with higher taxes and lower spending, reducing the level of output.

When expected inflation is too low, raising it is also hard and costly. Japan has been in this scenario for almost 20 years now. The Bank of Japan has tried forward guidance, quantitative easing, qualitative easing, yield curve control, and a series of other policies, all to no avail. It seems to require a great deal of commitment to convince economic agents to move from expecting 1% inflation in the long run to expecting 2% instead. Economic theories do not provide a clear answer for why it is so, but the Japanese experience gives credence to the fear that raising long-run expected inflation is a hard task.

As Mario Draghi stated unequivocally in Sintra in June 2018: "What is key is that inflation expectations remain well anchored".

3. Are long-run inflation expectations anchored at 2% today?

The most natural way to measure what economic agents expect inflation will be in the long run is to ask them. In the euro area, four times a year the ECB Survey of Professional Forecasters asks a select group of forecasters working for large firms what they expect inflation to be on average over the next five years. Between 2008 and 2018, the median response was never below 1.8%. It was never above 2.0%. Given the ECB's target for inflation of 2% or below, long-run expectations seem very well anchored.

In the United States, the Survey of Professional Forecasters run by the Federal Reserve Bank of Philadelphia asks a similar question with reference to the next five or 10 years, on average. Focusing on the five-year response, for comparability, again over the decade until 2018, the median answer was always between 2.1 and 2.4%. The Michigan survey asks a few hundred households every month in a rotating panel to report what they expect inflation to be in the long run. It is well known that the answers tend to be above actual inflation and are quite volatile, which is probably accounted for by households being not so well informed. Still, the range of answers over the 10 years before 2018 was only 2.5 to 3%. Again, US long-run inflation expectations seem remarkably well anchored.

Finally, for the United Kingdom, the Survey of Economic Forecasters asks professional forecasters what they expect inflation to be on average over the next three years. The range here was 1.9 to 2.2% in 2008–18.

Looking at these data alone, the answer is loud and clear: long-run inflation expectations seem very well anchored. Yet, I am sceptical, and dare I say, fearful. Household expectations of events quite far away are quite sluggish. People are inattentive, and they heavily discount the benefits of good forecasts far in the future relative to the costs of paying attention today. At the same time, when expectations do move, they do so persistently. The other side of the inattention is that there is a great deal of sticky information. It can take quite a while for the anchor to change in the expectations, but it will likewise take a lot for it to move back to the desired original anchor. Given the long and variable lags from monetary policy actions to inflation outcomes, by the time the surveys change, it is often too late for the central bank to do something about it. The pain of trying to shift back long-run inflation expectations that I described earlier in this lecture becomes inevitable.

Japan is a case in point when it comes to the inflation anchor moving down in a persistent way. The answers to the Consensus Economics survey of economic forecasters about long-run inflation were quite close to 2% between 1988 and 1996. By the end of the century, they had fallen to around 1%. Since then, they have rarely exceeded 1.5%.

4. A faster-moving alternative: market measures

An alternative measure of inflation expectations to that provided by surveys comes from financial market prices. Looking at the prices at which inflation swap contracts trade, or the difference between the yield on nominal government bonds and inflation-indexed government bonds, one can obtain some measure of the expectations of the participants in these financial markets. In the major financial markets, these financial contracts have been sufficiently liquid for about one decade that their prices reveal reliable information.

Looking at these measures over the same five-year horizon for the decade 2008–18 provides a much more sobering view. For the euro area, the standard deviation of long-run expected inflation was 0.5%, as it fluctuated between as high as 2.6% in Q1 2008 and as low as 0.6% in Q4 2014. The ratio of the standard deviation of expected long-run inflation to actual inflation is a strikingly high 0.44. For the United States, the standard deviation is also 0.5%, and again this is as high as half of the standard deviation of actual inflation. Expected long-run US inflation was as high as 3.2% in Q2 2008 and as low as 1.3% in Q2 2015.

Especially worrying are the numbers for 2019 so far. In both the euro area and the United States, expected long-run inflation has been steadily falling, and in the euro area it is already below 1.3%. At face value, these numbers suggest that long-run expected inflation may be about to fall below the central banks' targets. If so, the central banks that benefited from (and contributed to) a favourable anchoring of long-run inflation expectations around their target over the last 20 years may be about to endure the pain that comes with a change in that anchor.

Before making that inference it is important to ask what could lead to such a steep decline in market expectations contrary to survey expectations. One hypothesis is that inflation risk premia have declined, becoming negative and large in absolute value. Qualitatively, this story makes sense. Perhaps the dominant fear in 2019 is that we will have deflation combined with economic stagnation. The bad, high marginal utility state of the world that investors in financial markets want to insure against would be that of inflation being too low. Thus, inflation risk premia are negative, and a perceived higher likelihood of this deflation-stagnation state of the world has driven down market inflation expectations over the last 12 months.

While this hypothesis is sensible, it does not work quantitatively. In the inflation options market, we find the same traders who trade inflation swaps. The prices in these markets give a measure of the probability that inflation will be below –2% over the next five years on average. That number is small, but more importantly, it has moved very little in the last 12 months. If the deflation-stagnation fear was driving the fall in expected inflation, then it should show up significantly in the probability of the event. That probability would have risen from say 3% to 30% to account for the 1% change in the compensation for risk in expected inflation. Instead it changed by a couple of percentage points, far too modest to be consistent with raised fears of this state of the world.

More generally, historically, the standard deviation of expected long-run inflation according to the options moves little from quarter to quarter in spite of large movements in expected inflation. We would expect compensation for risk to come tightly associated with the perceived variance of inflation. In equity markets, the variance of expected equity returns in option markets often moves in excess of 10% within one year, and so do expected returns and the compensation for risk associated with them. In inflation markets, however, the standard deviation barely moves by more than 0.5% across years in the decade before 2018. This is less than one order of magnitude what is necessary to justify a compensation-for-risk interpretation of the movement in expected long-run inflation.

If not compensation for risk, then what is moving markets' long-run inflation expectations? Disagreement can do so, across two dimensions. First, disagreement between market participants and survey respondents, or, if you want, between the markets and the people. In the last 12 months, the markets have become significantly more pessimistic about future long-run inflation. In surveys of financial participants, where they are asked to report their subjective expectations, not their market-adjusted or risk-adjusted ones, the fall in expectations of inflation is clear. The public may be sluggish and inattentive, but these market traders are not, as they think about and trade on inflation information every day. The median response in these surveys has fallen significantly in 2019, explaining a large chunk of the decline in the market prices.

Second is disagreement among market participants – between the marginal trader (whose view the market price reflects) and the average trader. This shows up statistically as a change in the skewness of the distribution of the survey of inflation expectations among market dealers. In recent times, one sees the emergence of significant mass in the left tail of distributions of beliefs across traders. This shifts the marginal trader to the left of the distribution away from the median, explaining another part of the sharp fall in the market prices.

This alternative explanation is worrying for the anchoring of long-run inflation expectations around the target. Markets often lead people. A decline in the market perceptions may be a leading indicator that public perceptions are about to fall as well. That is, unless policy does something about it.

5. Policy actions to re-anchor expectations

The experience of the euro area in 2014–15 is instructive with regard to what policy can do when long-run inflation expectations start falling in a way that threatens a change in the anchor. During that time, there was a similar decline in market expectations to the one that we have seen in the last 12 months. We also saw a similar decline in the subjective belief in surveys of market participants. And finally, we also saw a similar change in the skewness of the distribution of expectations. By the middle of that period, there was a similar fear that the anchor for long-run inflation expectations was about to fall.

In that period, though, the ECB acted very aggressively. It implemented quantitative easing through its asset purchase programme, expanding the size of its balance sheet significantly, and buying and directly holding long-term government bonds. It further extended its period of forward guidance in the commitment to keep interest rates very low for a prolonged period of time. Its commitment to keeping inflation anchored at 2% was made clear and backed by expansionary policies when the anchor seemed to be falling. In 2016 and 2017, market expectations reversed track. Expected long-run inflation rose and went back to the 2% target.

The first lesson from this experience to central bankers is: be aggressive. When there is a fear that long-run inflation expectations are about to move, respond right away. Realise that a change in the anchor is one of the biggest dangers that a central bank can face.

Comparing this period with Japan in the late 1990s and early 2000s leads to the second lesson. The Bank of Japan was at the time still focused on restoring financial stability and dealing with the cleaning-up of banks and the associated outstanding bad credit. While it stated its commitment to 2% inflation, it gave the impression that it was happy to somewhat undershoot this target. By the time it adopted expansionary policies, it was already a few years since the decline in long-run inflation expectations. As soon as actual inflation started edging upwards, the Bank of Japan started discussing policy normalisation and reversing expansionary policies. It has since then been stuck with low long-run expected inflation.

Thus, the second lesson is: mean your commitment to the target, for it will be tested. In the case of the euro area today, inflation has been below 2% for almost five years now. By itself, the deviation each year has been relatively small, but once they are accumulated over these many periods, they imply that the euro area is now 6.8% below its target price level. At the same time, during this period and now, the real exchange rate has been appreciating between the core areas of the euro and the periphery regions. Therefore, correcting the deviation of the price level from target requires inflation to be well above 3% in Germany for a few years.

A third lesson is: do not be afraid of inflation going up to 3 or 4%. Policymakers today may worry about inflation being 1%, and want to raise it to 2%. But if they are mortally afraid that it may rise to 3%, then they are probably going to fail to achieve their goal. Every month a different shock is going to push inflation up or down by quite a few decimals. It is quite likely that on the path from 1% to 2%, a shock here and there will push inflation up to 3% or more. If policy reacts strongly to these and reverses course in an attempt to raise inflation expectations, then it will never get on the path towards a long-run inflation anchor of 2%.

Furthermore, shocks in the other direction will sometimes result in inflation staying at 1% or lower even as the central bank is doing all it can to raise it to 2%. If the anchor has indeed fallen below the 2% target, and the central bank is in the difficult position of pulling it back up, doing all it can to get the attention of private agents may be worth it. This may well include aiming for inflation temporarily above 2%.

An example comes from the United Kingdom in 2017–19. The effects of Brexit, and the loss in value of the pound through several moments of uncertainty, have led to inflation in the past two years being routinely around 3%. Perhaps it is not strange that, among the advanced economies, the United Kingdom is the one where long-run inflation expectations seem to be solidly anchored at 2% rather than trending down.

The fourth and final lesson is a familiar one to modern macroeconomics. Policy regimes, not isolated policies, are needed to sustain long-run outcomes. Most of the time, it makes sense to have monetary policy be set with an eye on inflation, and for the central bank to ignore the fiscal consequences of its actions. The separation between monetary and fiscal policy then implies that central banks refrain from engaging in operations that have too large a fiscal footprint. In exchange, they are independent from the fiscal authorities.

Sometimes, though, the monetary-fiscal separation can, and perhaps should, be broken. A tried and tested way to raise inflation and inflation expectations all the way into three digits is to give fiscal goals to monetary policymakers. Modest fiscal interventions by monetary policymakers directed to producing fiscal revenues that are transferred either to the government or directly to the public may well be able to raise long-run inflation expectations. In general, given our current knowledge, it is hard to calibrate these fiscal interventions to make inflation hit its target. Most likely, breaking the separation between fiscal and monetary policy will produce runaway inflation rather than slightly higher inflation as desired. However, keeping this option as an escape clause may play a role in keeping long-run inflation expectations from falling below target.

6. Communication and expectations

The goal of communication policies is ultimately to manage the expectations of economic agents. While communication that is not backed by fundamentals cannot accomplish much, at least in the long run, there is much work to be done by a central bank in explaining its policies and their goals. It is an essential part of what monetary policy must do, given the dependence of outcomes on agents' expectations. When it comes to the topic of this lecture, the anchoring of long-run inflation expectations, this becomes even more important. By communicating effectively, the central bank gains credibility with the public, and reveals its commitment to the targets. Every modern central bank today invests resources in communicating effectively and worries about the failures and successes of these messages.

Most of this existing communication is useful and especially important in light of keeping the anchor of long-run inflation expectations on target. Through communication, central banks have repeated what their target is, and reinforced their commitment to achieve it. Especially when it comes to unconventional policies, like those that involve the composition of assets in the balance sheet, communicating what the central bank is doing and why it is doing it has been fundamental for those policies to be able to affect expectations. When it comes to some policies, like forward guidance, that rely almost entirely on being able to shift expectations, then communication is in many ways what the whole policy is about.

Central banks do a worse job of communicating the links between policies and goals. Using reason, logic and, especially, economics, central banks must explain why they have used some tools given their targets. Explaining economics to the general public is hard, and central banks are not alone in not being successful. I am more worried, though, that in pursuing this worthwhile goal, they have overstepped.

A few central banks today go far beyond communicating goals, targets, and their links. They state that their goal is to "engage a broader cross-section of society". They worry that most people have no idea who the head of the central bank is right now, or that a vast majority is unaware of whatever the last communication was by the central bank. As a result, a few central bankers have started making regular speeches about topics that are more likely to get them onto the front page of newspapers. Climate change, trends in inequality of labour income, or changes in long-run business dynamism and competition are some examples. Invariably, the issues involved are important. Arguably, they matter more for social welfare than controlling inflation. There is therefore a good case to make for central banks to talk about them: they allow them to be relevant, as well as to focus on what matters to people and their well-being.

At the same time, there is very little that the central bank can do about these issues. Continuing with the focus on the long run, I started this lecture by stating why inflation may well be the sole objective for central banks in the long run. The arguments I made for why, maybe, neither real activity nor financial stability should be additional objectives, apply with much greater strength to inequality, competition, or the average temperature. Moreover, these topics are by their very nature controversial. Partly, this is precisely why they get so much media attention. It is almost impossible for central banks not to be dragged to these controversies. Being dragged into a controversial debate when you can do close to nothing to affect the debated outcomes does not seem like effective communication.

Another form of communication that central banks have been quick to embrace is simple messages to the public of the type: "trust me; I know what I am doing". Central banks have started producing video clips, cartoons, music videos, and different forms of media outreach where the message is so simplified that it boils down to bland statements that the central bank is very important and that inflation is very bad. Central bankers have been quite willing to support stories that they saved the world during the financial crisis, and/or have prevented more than one recession through their diligent actions. In many ways, this is fine, and appropriate. But it also implies that when a recession or a financial crisis comes, or even when inflation deviates from target for a few years, central banks will be blamed. After all, they communicated clearly that the absence of these bad outcomes was to their credit. And yet, each of these outcomes will inevitably happen given the limits of what monetary policy can actually achieve.

My worry is that being relevant and simple may be attractive but it will backfire. It may erode the trust that the public has in the central bank; trust which will be especially important if the next challenge for the central bank is to raise long-run inflation expectations back to target.

7. Conclusion

This lecture asked a few questions and provided answers along the way. The conclusion section is a good place to restate them in a shortened version (albeit a less nuanced one than is adequate):

- What is the long-run goal of the central bank? Low and stable inflation, alone.
- Why are long-run inflation expectations central? Because anchoring inflation is anchoring expectations.
- Are long-run inflation expectations anchored on target today? Surveys make it seem so, but they are too sluggish to allow detection of incipient changes.
- Do markets provide a better measure? Yes; they are more forward-looking, and show that there is cause for concern in the euro area and United States.
- What policies can re-anchor expectations on target? Be aggressive; mean it; don't fear inflation at 3–4%; fiscal escape clauses matter.
- What is the role of communication? Key to describe goal, tools and their links, but central banks need to be careful when trying to be relevant and simple, as this can lose them trust along the way.

A final word directed to the emerging economies represented at this conference. The reduction in long-run inflation expectations in European countries, the United States, the United Kingdom and Canada throughout the 1980s had persistent effects that spread to the rest of the world. Most central banks since then have adopted the tools and approaches followed by those central banks in terms of institutional design for independence, adoption of numerical inflation targets, operational procedures for setting interest rates, and the like. If low long-run inflation expectations turn out to be the new challenge for the next few years, and those same central banks find ways to raise these expectations, sooner or later this is likely to have an impact on emerging economies as well.

What drives inflation in advanced and emerging market economies?

Güneş Kamber, Madhusudan Mohanty and James Morley[1]

Abstract

Kamber et al (2020) investigate possible changes in the driving forces of inflation for a panel of 47 advanced and emerging market economies over a sample period from 1996 to 2018. Overall, the results support an open economy hybrid Phillips curve model of inflation with increased weight on expected future inflation and an important role of the foreign output gap. The estimated effects of inflation expectations, output gaps, exchange rate pass-through and oil prices are heterogeneous across different economies, with generally larger effects of external driving forces for emerging market economies. Also, despite some structural changes, the parameters of the model show a surprising degree of stability before, during and after the Great Financial Crisis. For many economies, the estimated effect of a given variable does not change at all over the full sample period, while the behaviour of the variables in the model can explain patterns of changes in both the level and volatility of inflation over time.

Keywords: open economy Phillips curve; structural breaks; inflation expectations; exchange rate pass-through; inflation volatility.

JEL classifications: E31, F31, F41.

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1. Introduction

Should we be surprised by the behaviour of inflation across different economies in recent years? This paper summarises our recent work (Kamber et al (2020)), which investigates possible changes in the driving forces of inflation for a panel of 47 advanced and emerging market economies over a sample period from 1996 to 2018 that includes the Great Financial Crisis (GFC) in the late 2000s. Our results suggest that we should not be surprised by the behaviour of inflation, as it is consistent with an open economy hybrid Phillips curve model of inflation for which we find an increased weight on expected future inflation and an important role of the foreign output gap.

There is an enormous literature on the behaviour of inflation, with many recent studies focusing on the relative stability of inflation worldwide in the face of the GFC and record low interest rates in its aftermath. This stability has been attributed to a flattening of the Phillips curve and an anchoring of inflation expectations (eg IMF (2013)). The idea that the relationship between inflation and the real economy would change given a different policy environment is the prime example of the original Lucas (1976) critique, with anchored inflation expectations and constraints on monetary policy due to the effective lower bound on interest rates clearly corresponding to large changes in the policy environment for central banks in recent years. Another hypothesised source of change in inflation drivers in recent years is increased foreign competition and trade integration (eg Forbes (2018)).

When we consider the behaviour of inflation for a number of advanced and emerging market economies before, during and after the GFC, we find that it is well captured by an open economy hybrid Phillips curve model that includes backward- and forward-looking inflation expectations, domestic and foreign output gaps, exchange rate pass-through and oil prices. Our results support an increased weight on expected future inflation and an important role of the foreign output gap. Perhaps surprisingly though, at least given the Lucas critique, structural break tests suggest that most model parameters are stable for a majority of economies throughout the sample period, including during the crisis. We find widespread importance of all of the potential driving forces in the model, but their effects are quite heterogeneous across different economies, with the external driving forces generally having larger effects for emerging market economies. Notably, we find that the behaviour of the variables in the model can explain patterns of changes in both the level and volatility of inflation over time.

Our paper relates to a number of strands of literature on what drives inflation, including on differences for advanced and emerging market economies (eg Blanchard et al (2015), IMF (2016, 2018), Miles et al (2017), Jorda and Nechio (2018), Ha et al (2019), Kamber and Wong (2020)), the role of expectations (eg Fuhrer (2012), Coibion and Gorodnichenko (2015), Cecchetti et al (2017), Ball and Mazumder (2019)), global influences (eg Borio and Filardo (2007), Monacelli and Sala (2007), Ciccarelli and Mojon (2010), Guerrieri et al (2010), Ihrig et al (2010), Milani (2010), Mumtaz and Surico (2012), Bianchi and Civelli (2015), Auer et al (2017)) and exchange rate pass-through (eg Choudhri and Hakura (2006), Mihaljek and Klau (2008), Jasova et al (2016)).

The rest of this summary is organised as follows. Section 2 describes our empirical model and the data. Section 3 summarises the key empirical results. Section 4 concludes.

2. Model and data

For each economy i, we consider an open economy hybrid Phillips curve specification for inflation:

$$\pi_{it} = \beta_{0i} + \beta_{1i}\pi_{it}^{e,bwd} + \beta_{2i}\pi_{it}^{e,fwd} + \beta_{3i}\tilde{y}_{it} + \beta_{4i}\tilde{y}_{it}^{*} + \beta_{5i}\Delta_4 e_{it} + \beta_{6i}\Delta_4 p_{t-1}^{oil} + \varepsilon_{it}, \quad (1)$$

where π_{it} is a quarterly measure of year-on-year inflation, $\pi_{it}^{e,bwd}$ is a measure of backward-looking inflation expectations, $\pi_{it}^{e,fwd}$ is a measure of forward-looking inflation expectations, $\tilde{y}_{it}$ is a measure of the domestic output gap, $\tilde{y}_{it}^{*}$ is a measure of the foreign output gap, $\Delta_4 e_{it}$ is a quarterly measure of the year-on-year change in the exchange rate in per cent, $\Delta_4 p_{t-1}^{oil}$ is the lagged quarterly measure of the year-on-year change in world oil prices in per cent (measured in US dollars), and ε_{it} is a residual inflation shock that is possibly heteroskedastic with assumed scale-equivariant long-run variance $\bar{\sigma}_{\varepsilon i}^{2}$.

To reduce the number of independent parameters and increase the precision of our estimates, we also consider a restriction in estimation that the coefficients on backward- and forward-looking inflation expectations are weights that sum to one: $\beta_{2i} = 1 - \beta_{1i}$. This restriction is imposed by considering the following regression:

$$\pi_{it} - \pi_{it}^{e,fwd} = \beta_{0i} + \beta_{1i}(\pi_{it}^{e,bwd} - \pi_{it}^{e,fwd}) + \beta_{3i}\tilde{y}_{it} + \beta_{4i}\tilde{y}_{it}^{*} + \beta_{5i}\Delta_4 e_{it} + \beta_{6i}\Delta_4 p_{t-1}^{oil} + \varepsilon_{it}.$$

Notably, even if inflation and inflation expectations are non-stationary for some of the economies under consideration, imposing this restriction serves to render all of the variables in the regression stationary given cointegration between inflation and inflation expectations such that expectation errors are $I(0)$. This is important because stationarity is often a maintained assumption when conducting structural break analysis with unknown break dates (Bai and Perron (1998, 2003), Qu and Perron (2007)), which is part of our analysis. Assuming the restriction is valid, it is also useful for econometric identification of structural breaks at unknown break dates to have fewer independent parameters in our test regressions.

For estimation, we follow much of the empirical literature and assume that the explanatory variables in equation (1) are exogenous or at least predetermined in the sense that the inflation shock ε_{it} only affects them with a lag. Also, we assume no serial correlation in ε_{it}, an assumption that is supported in practice by the inclusion of lagged inflation π_{it-1} in our regressions as the measure of backward-looking inflation expectations.[2] Any violation of these assumptions would result in biased and inconsistent estimates. Reassuringly, though, we find that our results are generally robust, albeit not as statistically significant, when considering lagged measures of explanatory variables, which are predetermined by construction.[3]

We construct a balanced panel data set of the variables in equation (1) for 47 economies over a sample period of Q1 1996 to Q3 2018. For inflation, we use the

2 The inclusion of lagged inflation also means that equation (1) can be thought of as an autoregressive distributed lag model, at least if the other variables are assumed to be exogenous, where the long-run effects of shocks to the other variables would be equal to the short-run effects multiplied by $1/(1 - \beta_{1i})$.

3 We include lagged oil prices in our main specification because we want to control for the possibility that inflation shocks for some large economies could feed into contemporaneous changes in oil prices. However, results are also robust to considering contemporaneous oil prices.

four-quarter change in 100 times the log of headline CPI obtained from national data sources and the BIS. For our measure of backward-looking inflation expectations, we use lagged inflation, as noted above and as is standard in the literature. For our measure of forward-looking inflation expectations, we use the one-year-ahead survey forecasts obtained from Consensus Economics and the BIS, where the short-term horizon for expectations is consistent with the standard representation of the New Keynesian Phillips curve and allows for consideration of a much broader set of economies than would be possible for longer-term inflation expectations.[4] We note that, although the survey forecasts may have some backward-looking element to their formation, the coefficient on expected future inflation should capture the impact of the forward-looking element of the survey forecasts given the control for lagged inflation in our regressions. For measuring the output gaps, we use the HP filter (λ = 1,600) applied to seasonally adjusted quarterly log real GDP obtained from national data sources, the IMF and the BIS.[5] Following Borio and Filardo (2007), the foreign output gap is constructed using trade weights for the 10 largest trading partners obtained from UN Comtrade and the BIS, with weights updated annually.[6] For the exchange rate, we use a nominal effective exchange rate index obtained from Bruegel Datasets. For oil prices, we use the WTI index obtained from Datastream.

Based on BIS classification, the 47 economies include 20 advanced economies (AEs) and 27 emerging market economies (EMEs). By contrast, based on IMF classification, the same 47 economies consist of 30 AEs and 17 EMEs. Therefore, using standard two-letter codes for the 47 economies, we consider the following three groups:[7]

1. AT, BE, CH, DE, DK, ES, FI, FR, GB, IE, IT, NL, NO, PT, SE, AU, CA, JP, NZ, US
2. CZ, EE, GR, LT, LV, SI, SK, KR, HK, SG
3. HU, PL, RU, TR, ZA, CN, ID, IN, MY, PH, TH, AR, BR, CL, CO, MX, PE

The BIS classifies the economies in the first group as AEs and those in the second and third groups as EMEs, while the IMF classifies the economies in the first and second groups as AEs and those in the third group as EMEs. We consider robustness of our results for AEs and EMEs to both classifications.

4 Monthly surveys on year-end forecasts for the current and following years are converted to fixed-horizon 12-month-ahead forecasts following Yetman (2018). Expected future inflation is measured in terms of the expected change in 100 times the log CPI over the next four quarters.

5 For robustness we also consider an alternative approach to trend-cycle decomposition based on the BN filter with dynamic demeaning that Kamber et al (2018) show provides more reliable real-time estimates of the output gap than the HP filter. The results are largely robust, suggesting that real-time issues are less important for examining historical driving forces of inflation than they would be for understanding current inflation pressures.

6 The trade weights for a given economy and trading partner are defined as the ratio of trade openness between the economy and trading partner (exports plus imports) divided by the total trade openness of the given economy. We consider the 10 largest trading partners in the data set.

7 The listed order for each group is alphabetical within each region, ie Europe (including ZA), Asia-Pacific and North America (including CA and US), and Latin America, respectively.

3. Empirical results

3.1 Panel estimates

We first consider panel estimates to compare to Forbes (2018). The estimation assumes the slope coefficients in equation (1) are the same across different economies (ie $\beta_{ji} = \beta_j$ for $j = 1,\ldots,6$). As in Forbes (2018), we allow for random effects and a common structural break in slope coefficients in Q4 2006 by interacting explanatory variables with a dummy variable $D_t = 1$ for $t \geq$ Q1 2007 and 0 otherwise.

Table 1 summarises the results in Kamber et al (2020) for the panel estimation. Despite some differences in the data and model specification, the estimates are similar to those in Forbes (2018).[8] All of the variables in equation (1) appear to be significant driving forces of inflation. There is a higher estimated weight on lagged inflation than expected future inflation, although this result was somewhat ambiguous in Forbes (2018) depending on the measure of inflation considered. In our unrestricted estimation, the weight on expected future inflation increased after the onset of the crisis. In our restricted estimation, the slope of the domestic Phillips curve flattened, although this is not significant (nor was it for headline CPI in Forbes (2018)). In both cases, the slope of the foreign Phillips curve increased, while the degree of exchange rate pass-through decreased, significantly so in the unrestricted case, while the decrease was not significant for Forbes (2018).

We highlight that the unrestricted and restricted estimates are quite similar, justifying the imposition of the theoretically motivated restriction that the coefficients on lagged inflation and expected future inflation correspond to weights that sum to one. Although the estimates are similar, there is one noticeable difference in terms of the implied change in the weight on expected future inflation with the onset of the crisis. In the restricted case, there is basically no estimated change in the weight. This difference in inference presumably reflects the fact that the inflation data contain relatively more information about the relationship between lagged inflation and current inflation than about the relationship between expected future inflation and current inflation.

So, to summarise the more robust results from the panel estimation, it appears that both domestic and foreign driving forces of inflation are important, with the restriction that the coefficients on lagged inflation and expected future inflation correspond to weights that sum to one supported by the data. The estimates suggest an increase in the slope of the foreign Phillips curve after the onset of the crisis. Less clear is whether the weight on expected future inflation increased or the slope of the domestic Phillips curve decreased after the onset of the crisis. But overall the estimates are in line with those from the similar panel analysis in Forbes (2018).

[8] Forbes (2018) considers a similar, but not identical, set of economies and sample period. In her baseline specification, she also considers quarterly inflation, both headline and core, rather than year-on-year for the dependent variable, although she considers lagged year-on-year inflation, as we do, to measure backward-looking expectations. She uses a five-year-ahead survey forecast to measure forward-looking inflation expectations, different measures of domestic and foreign output gaps, the change in per cent in the real effective exchange rate over a two-year horizon, and the lagged change in per cent in world oil prices over a one-quarter horizon. She also includes commodity prices and a measure of world producer price dispersion in her baseline model.

Panel coefficient estimates Table 1

Variable	No break		Break	
	Unrestricted	Restricted	Unrestricted	Restricted
$\pi_{it}^{e,bwd}$	0.584***	0.583***	0.588***	0.580***
	(0.045)	(0.044)	(0.051)	(0.049)
$\pi_{it}^{e,fwd}$	0.414***	0.417***	0.391***	0.420***
	(0.056)	(0.044)	(0.058)	(0.049)
$\tilde{y}_{it}$	0.074**	0.075**	0.138**	0.136**
	(0.035)	(0.037)	(0.068)	(0.067)
$\tilde{y}_{it}^{*}$	0.099**	0.097**	–0.118**	–0.082
	(0.041)	(0.045)	(0.055)	(0.061)
$\Delta_4 e_{it}$	–0.085***	–0.085***	–0.109***	–0.098***
	(0.024)	(0.023)	(0.030)	(0.027)
$\Delta_4 p_{t-1}^{oil}$	0.005***	0.005***	0.005***	0.003
	(0.001)	(0.001)	(0.002)	(0.002)
$D_t \cdot \pi_{it}^{e,bwd}$			–0.001	–0.022
			(0.040)	(0.045)
$D_t \cdot \pi_{it}^{e,fwd}$			0.102***	–0.022
			(0.036)	(0.045)
$D_t \cdot \tilde{y}_{it}$			–0.118	–0.113
			(0.074)	(0.078)
$D_t \cdot \tilde{y}_{it}^{*}$			0.257***	0.260***
			(0.069)	(0.072)
$D_t \cdot \Delta_4 e_{it}$			0.071**	0.047
			(0.029)	(0.029)
$D_t \cdot \Delta_4 p_{t-1}^{oil}$			0.001	0.003
			(0.002)	(0.002)

For panel estimation, we assume $\beta_{ji} = \beta_j$ for all i, but we allow for economy-specific random effects. For the restricted case, estimates for both β_1 and β_2 are reported using $\beta_2 = 1 - \beta_1$. The dummy variable $D_t = 1$ for $t \geq$ Q1 2007 and 0 otherwise allows for a structural break in slope coefficients β_{ji} for $j - 1, \ldots, 6$ with the onset of the GFC and changes in slopes are reported. Robust standard errors with clustering by economy are reported in parentheses and significance is based on two-tailed t tests (***significant at 1%, **significant at 5%, *significant at 10%).

3.2 Economy-by-economy estimates

One issue with the panel estimates in Table 1 is that they do not allow for heterogeneity in the slope coefficients in equation (1) or in the existence and timing of structural changes in the slope coefficients. We find that allowing for these forms of heterogeneity is critically important for inferences about the effects of different potential driving forces of inflation over time.

First, we examine slope coefficient heterogeneity by testing whether estimated coefficients from an economy-by-economy estimation of equation (1) are significantly different to the estimated coefficients from the panel estimation:

$H_0: \beta_{ji} = \hat{\beta}_j^{panel}$.[9] Table 2 summarises the results in Kamber et al (2020) for this test. We are able to reject the null hypothesis far more than the 5% test size for each slope coefficient and for as many as 74% of the economies in the case of the coefficient on the exchange rate. Thus, we can conclude that differences in estimates are not simply due to sampling error, but there is true underlying heterogeneity in effects of the various potential driving forces of inflation across different economies.

Rejection rates in support of coefficient heterogeneity — Table 2

Variable	$H_0: \beta_{ji} = \hat{\beta}_j^{panel}$
$\pi_{it}^{e,fwd}$	26%
$\tilde{y}_{it}$	21%
$\tilde{y}_{it}^{*}$	15%
$\Delta_4 e_{it}$	74%
$\Delta_4 p_{t-1}^{oil}$	21%

We consider a two-tailed t test using HAC standard errors calculated according to Andrews and Monahan (1992).

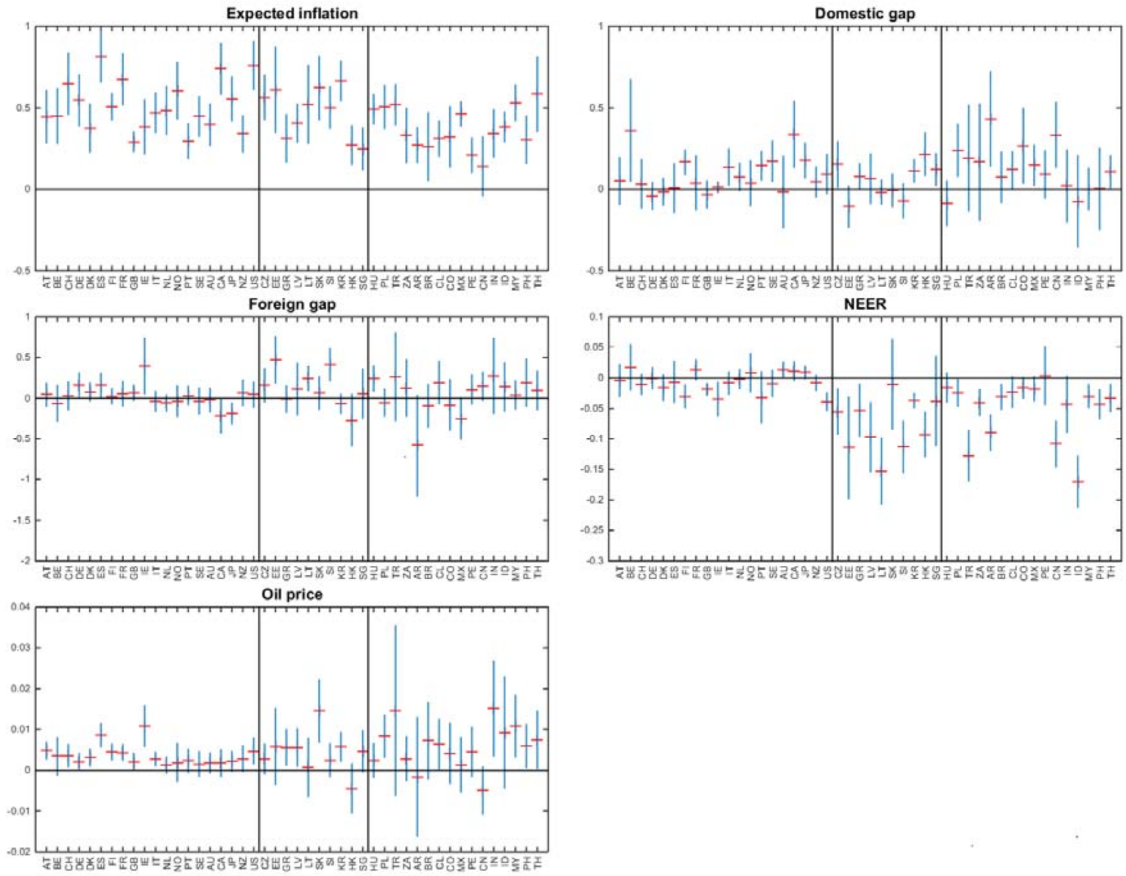
Graph 1 displays the economy-by-economy coefficient estimates and 95% confidence bands from Kamber et al (2020), with the estimates sorted by the three different groups of economies listed in the previous section. Interestingly, the results for expected inflation and the domestic output gap suggest that the second group of economies is more like the first group (that are always classified as AEs) than the third group (that are always classified as EMEs), consistent with the IMF classification of economies, while the results for the foreign output gap, the exchange rate, and oil prices suggest that the second group is more like the third group (always classified as EMEs) than the first group (always classified as AEs), consistent with the BIS classification of economies. However, the key result is that there is considerable heterogeneity in the estimated effects of the potential driving forces of inflation, with the external driving forces having typically larger effects for EMEs than AEs.[10] The weight on expected future inflation is generally higher for AEs than EMEs, while the slope of the domestic Phillips curve is just heterogeneous, with no clear patterns across AEs and EMEs.

9 For the economy-by-economy analysis, we always consider the restricted case that imposes that the coefficients on lagged inflation and expected future inflation are weights that sum to one. Thus, for simplicity, we do not report results for the coefficient on lagged inflation, which are directly implied by the results for the coefficient on expected future inflation.

10 Also consistent with heterogeneity, Kamber et al (2020) find that the mean effects are noticeably larger in magnitude than median effects for the domestic output gap, the exchange rate, and oil prices, suggesting that there is a small subset of economies with particularly large sensitivities to these variables.

Economy-by-economy coefficient estimates

Graph 1



Results are reported for slope coefficients on expected future inflation, the domestic output gap, the foreign output gap, the nominal effective exchange rate, and oil prices, with two-letter codes for different economies listed on the x-axis (RU is excluded given the much larger scale for its confidence interval). Point estimates are red crosses and confidence intervals are blue lines. Confidence intervals are based on inverted t tests using HAC standard errors calculated according to Andrews and Monahan (1992).

Next, we examine heterogeneity in the existence and timing of structural changes in the slope coefficients. To do so, we apply tests for possible multiple structural breaks in coefficients and/or the error variance in equation (1) at unknown break dates according to the procedures in Qu and Perron (2007).[11] Given the existence of structural breaks, we test whether specific coefficients change using likelihood ratio tests for $H_0\colon \beta_{jit} = \beta_{ji}$. Table 3 summarises the results in Kamber et al (2020) based on these tests. First, we find around two structural breaks in the model parameters on average. However, this appears to correspond mostly to changes in the long-run variance of inflation shocks. Notably, the median number of breaks for the slope coefficients on all but the foreign output gap is zero, while it is only one for the foreign

[11] Tests and estimation of structural breaks allow for heteroskedasticity and serial correlation in the residuals. As is standard when estimating structural breaks at unknown break dates, we use trimming that restricts the minimum length between breaks to be at least 15% of the total sample period. Structural breaks are estimated sequentially. In principle, we allow up to five breaks. However, given some short subsample periods with sequential testing, we are often restricted to fewer possible breaks in practice. Fortunately, the maximum allowable number of breaks is almost never binding for the estimated number of breaks in slope coefficients.

output gap. Yet there is also clear heterogeneity regarding structural breaks, as there is at least one break in a given slope coefficient for a significant number (at least one quarter) of the economies under consideration.

Distribution of estimated number of breaks across different economies — Table 3

Variable	Mean	25th percentile	Median	75th percentile
$\pi_{it}^{e,fwd}$	0.47	0	0	1
$\tilde{y}_{it}$	0.62	0	0	1
$\tilde{y}_{it}^{*}$	0.66	0	1	1
$\Delta_4 e_{it}$	0.64	0	0	1
$\Delta_4 p_{t-1}^{oil}$	0.64	0	0	1
Total (incl LR variance)	1.70	1	2	2
Max allowed	2.74	2	3	4

The presence of structural breaks and whether they apply to a given slope coefficient is tested for based on Qu and Perron (2007) procedures.

Average estimated effects — Table 4

	No break		Break	
Variable	Panel	Heterogeneity	Panel	Heterogeneity
$\pi_{it}^{e,fwd}$	0.417	0.456	0.408	0.528
$\tilde{y}_{it}$	0.075	0.103	0.077	0.068
$\tilde{y}_{it}^{*}$	0.097	0.047	0.054	0.076
$\Delta_4 e_{it}$	–0.085	–0.041	–0.073	–0.036
$\Delta_4 p_{t-1}^{oil}$	0.005	0.005	0.004	0.004

The average effects are $\frac{1}{T}\sum_t \hat{\beta}_{jt}$ for panel estimation and $\frac{1}{NT}\sum_i \sum_t \hat{\beta}_{jit}$ for economy-by-economy estimation that allows for heterogeneity in slope coefficients.

In the cases where there is evidence of structural breaks, they are estimated to have occurred at different times for different economies.[12] The main relevance of this result is that accounting for breaks has a more notable impact on average estimated effects of the variables in the model than implied by just allowing for one break in Q4 2006 with the panel estimation. Table 4 illustrates this by summarising the average effects for the panel and economy-by-economy estimates from Kamber et al (2020), first when not allowing for breaks and then when allowing for breaks. The panel estimates are mostly similar in the two cases, while the average weight on expected future inflation and the coefficient on the foreign output gap are noticeably larger when allowing for breaks in the economy-by-economy analysis.

[12] Notably, estimated break dates occur throughout the trimmed sample period, with no particular pattern of clustering of breaks around the crisis years. Also, 95% confidence sets for break dates based on inverted likelihood ratio tests following Eo and Morley (2015) often exclude the crisis years.

Average estimated effects before and after the GFC by different groups of economies

Table 5

Variable	Economies	Q2 1996–Q4 2006		Q1 2007–Q3 2018	
		Classifications			
		BIS	IMF	BIS	IMF
$\pi_{it}^{e,fwd}$	All	0.511	0.511	0.543	0.543
	AEs	0.569	0.561	0.608	0.574
	EMEs	0.468	0.422	0.495	0.490
	Asian EMEs	0.424	0.404	0.550	0.561
	Latin America	0.369	0.369	0.380	0.380
$\tilde{y}_{it}$	All	0.048	0.048	0.088	0.088
	AEs	0.025	0.023	0.150	0.102
	EMEs	0.064	0.091	0.041	0.062
	Asian EMEs	0.052	0.016	0.085	0.072
	Latin America	0.114	0.114	0.079	0.079
$\tilde{y}_{it}^{*}$	All	0.087	0.087	0.067	0.067
	AEs	0.047	0.116	–0.077	–0.001
	EMEs	0.116	0.035	0.173	0.186
	Asian EMEs	0.127	0.219	0.115	0.269
	Latin America	–0.182	–0.182	0.133	0.133
$\Delta_4 e_{it}$	All	–0.037	–0.037	–0.035	–0.035
	AEs	–0.008	–0.025	–0.015	–0.030
	EMEs	–0.059	–0.060	–0.050	–0.043
	Asian EMEs	–0.061	–0.086	–0.041	–0.048
	Latin America	–0.016	–0.016	–0.036	–0.036
$\Delta_4 p_{t-1}^{oil}$	All	0.004	0.004	0.004	0.004
	AEs	0.003	0.003	0.004	0.005
	EMEs	0.004	0.005	0.004	0.004
	Asian EMEs	0.006	0.007	0.004	0.004
	Latin America	0.005	0.005	0.001	0.001

The average effects are $\frac{1}{NT}\sum_i \sum_t \hat{\beta}_{jit}$ for the economy-by-economy estimation.

Table 5 summarises the average estimated effects from Kamber et al (2020) for different groups of economies before and after the GFC. Expected future inflation has become a more important driving force of current inflation since the crisis for all but Latin American economies, with average weights increasing by a large amount for Asian economies and reaching above 60% for AEs. The average slope of domestic Phillips curves increased for AEs and Asian economies, while it decreased for Latin America. The average slope of foreign Phillips curves increased for EMEs, while it decreased to close to zero for AEs. There was a decrease in average exchange rate pass-through for EMEs, except for Latin America, which had an increase, while average pass-through remains low before and after the crisis for AEs. The average effect of oil prices slightly increased for AEs and decreased for EMEs. Thus, heterogeneity in structural changes across regions that is obscured in the panel estimation is clearly very important.

It is worth noting that allowing for structural breaks also has strong implications for the general relevance of the various potential driving forces of inflation. Table 6 reports rejection rates for whether slope coefficients are always equal to zero. When allowing for breaks, the various potential driving forces appear relevant for at least two thirds of the economies, with particularly notable increases in the apparent relevance of the external driving forces compared to the case of not allowing for breaks.

Rejection rates in support of variable relevance

Table 6

Variable	$H_0\colon \beta_{ji} = 0$	$H_0\colon \beta_{jit} = 0, \forall t$
$\pi_{it}^{e,fwd}$	26%	26%
$\tilde{y}_{it}$	21%	21%
$\tilde{y}_{it}^{*}$	15%	15%
$\Delta_4 e_{it}$	74%	74%
$\Delta_4 p_{t-1}^{oil}$	21%	21%

For the first hypothesis, we consider a two-tailed t test using HAC standard errors calculated according to Andrews and Monahan (1992). For the second hypothesis, we consider if we can either reject the first hypothesis or a likelihood ratio test for a structural break in a given parameter based on Qu and Perron (2007) procedures.

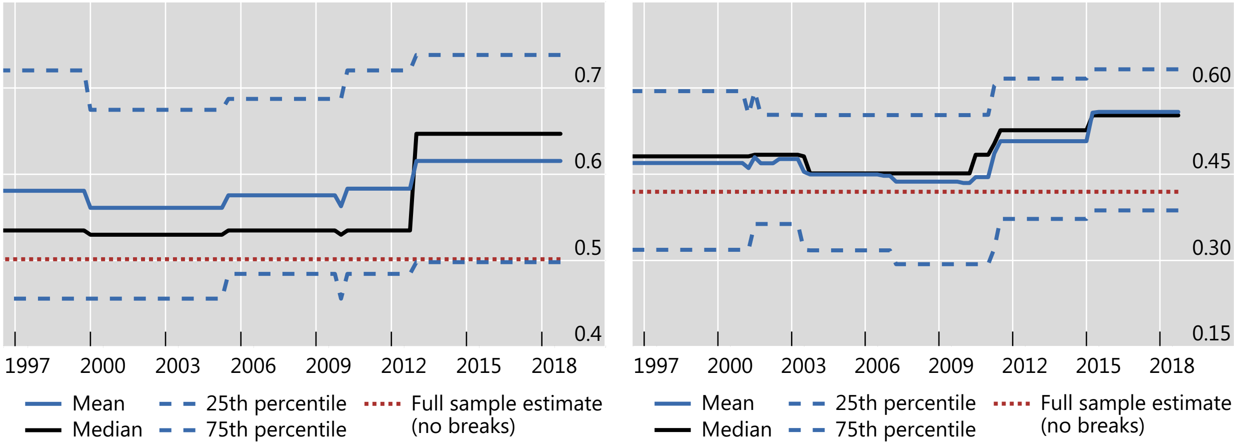
Thus, we can conclude that, when allowing for coefficient heterogeneity and heterogeneity in the existence and timing of structural breaks, the variables in an open economy hybrid Phillips curve model of inflation are relevant for a majority of advanced and emerging market economies. Furthermore, while the effects of the variables in the model are mostly stable for most economies, there are important changes for some economies that are quite heterogeneous by region. Perhaps the most notable finding when allowing for heterogeneity in structural breaks is for a high and increasing weight on expected future inflation for many advanced and emerging market economies. Graph 2 displays this result from Kamber et al (2020) by plotting the mean and percentiles of the distributions for both advanced and emerging market economies of estimated weights over the sample period conditional on estimated structural breaks. Allowing for breaks, the mean and median effects are higher than the average full-sample estimate and rising over time in both cases.

Role of expected future inflation for advanced and emerging market economies

Graph 2

Distribution of estimated coefficient on expected inflation, AEs

Distribution of estimated coefficient on expected inflation, EMEs



The estimated coefficients are conditioned on estimated break dates for structural breaks determined by Qu and Perron (2007) procedures. The mean estimate for groups of economies in the case of no breaks is also reported for comparison. Economies are grouped as AEs or EMEs according to the BIS classification.

3.3 What explains changes in the level and volatility of inflation?

Given that most slope coefficients are stable for a majority of economies throughout the sample period, we examine the extent to which the behaviour of the variables in the model can explain patterns of changes in the level and volatility of inflation over time.

First, we consider the presence of structural breaks in the long-run mean and variance of each possible driving variable and see if they are related to changes in the long-run mean and variance of inflation. Looking across all economies, we find that the average estimate of the long-run mean of inflation fell dramatically in the early 2000s from around to 7.5% to close to 3% by the end of the sample period, although the median estimate was more stable, but also fell, from just over 3% to just over 2%. Consistent with the large estimated weights on expected future inflation in the open economy hybrid Phillips curve model, structural break analysis suggests there were similar changes in the long-run mean of expected future inflation, implying that inflation expectations provide at least a proximate cause for the fall in actual inflation. There was also a corresponding reduction in average exchange rate depreciation across the full panel of economies that is consistent with long-run purchasing power parity and the convergence of long-run inflation rates. Likewise, the open economy hybrid Phillips curve model and apparent structural breaks in the long-run variance of expected future inflation and output gaps are consistent with a decline in average estimate of the long-run variance of inflation early in the sample period after the end of the Asian crisis and an increase during the crisis years and decrease afterwards. Notably, the changes in the long-run variance of inflation do not appear to be driven solely by structural breaks in the long-run variance of inflation shocks from the open economy hybrid Phillips curve model.

Next, related to structural breaks in the driving variables, we consider variance decomposition results based on the open economy hybrid Phillips curve model. Because the driving variables may be correlated with each other, we construct a model-implied proxy variance that ignores any such correlation:

$$\tilde{\sigma}^2_{\pi,it} \equiv \sum_j \hat{\beta}^2_{jit}\,\hat{\bar{\sigma}}^2_{j,it} + \hat{\bar{\sigma}}^2_{\varepsilon,it},$$

where $\hat{\beta}_{jit}$ is the estimated slope coefficient on variables j for economy i conditional on weighted average estimated structural breaks, $\hat{\bar{\sigma}}^2_{j,it}$ are the weighted average estimated long-run variances of the driving variables, and $\hat{\bar{\sigma}}^2_{\varepsilon,it}$ is the weighted average estimated long-run variance of the inflation shock. The weighted average estimates are based on confidence sets for structural break dates following the inverted likelihood ratio method in Eo and Morley (2015). Crucially, changes in the model-implied proxy variance $\tilde{\sigma}^2_{\pi,it}$ track changes in the weighted average estimated long-run variance of inflation $\hat{\bar{\sigma}}^2_{\pi,it}$ quite well over time, albeit with a fairly constant downward bias.

The main result for the variance decomposition analysis is that the combination of forward- and backward-looking inflation expectations explain a very large portion of inflation variation over time, with expected future inflation generally explaining a higher share for AEs than EMEs. Domestic and foreign output gaps explain relatively little of the overall variation in inflation at the beginning and end of the sample period, but they explain much more, especially for AEs, during the crisis years. The role of exchange rate pass-through declined in the 2000s, but increased again, especially for EMEs, in the 2010s. The role of oil prices also declined in the 2000s and then rose again in the 2010s, but they only ever explain a small share of the variance of inflation (less than 10% on average), despite the fact that we consider headline inflation.

4. Conclusions

In Kamber et al (2020), we find that an open economy hybrid Phillips curve provides a surprisingly stable structure for understanding the behaviour of inflation in advanced and emerging market economies in recent years. The result is surprising when one considers the Lucas critique along with the dramatic change in policy environment that occurred after the onset of the GFC. Our results do suggest an increase in the weight on expected future inflation in driving current inflation for both advanced and emerging market economies after the crisis. However, as would be predicted by a relatively stable open economy hybrid Phillips curve model, structural changes in expected future inflation, domestic and foreign output gaps, and, to a lesser extent, exchange rates and oil prices can explain patterns or changes in both the level and volatility of inflation throughout the sample period from 1996 to 2018.

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Comments on "What drives inflation in advanced and emerging market economies?"

Hans Genberg[1]

The paper by Kamber, Mohanty and Morley (KMM hereafter) is a carefully executed empirical investigation of the determinants of inflation in a broad sample of advanced and emerging market economies.

A number of recent multi-country empirical studies of inflation dynamics have used estimated Phillips curve relationships to highlight changes in parameters following the Great Financial Crisis (GFC).

KMM challenge previous findings based on panel regression methods, arguing that it is critically important to account for heterogeneity across economies and over time in assessing the importance of each determinant of inflation. They suggest that key findings in the literature using panel regression estimates are not robust. The reason is heterogeneity across economies and, to some extent, the use, in previous studies, of a single breakpoint in 2007.

The principal question underlying the empirical work is whether the inflation process has changed since the GFC. Answers are provided by estimates of the coefficients in a relatively conventional Phillips curve model of the inflation process:

$$\pi_{it} = \beta_{0i} + \beta_{1i}\,\pi_{it-1} + \beta_{2i}\,\pi_{it}^{e} + \beta_{3i}\,\tilde{y}_{it} + \beta_{4i}\,\tilde{y}_{it}^{*} + \beta_{5i}\,\Delta_4 e_{it} + \beta_{6i}\,\Delta_4 p_{it}^{oil} + \varepsilon_{it}.$$

A number of time dummies interacted with the independent variables are added to investigate possible changes in the value of the β coefficients. The i subscripts indicate that coefficients may be different across countries.

The specific questions addressed in the paper relate to the stability of the coefficients (the β:s) over time; the relative importance of forward-looking (expected) relative to backward-looking inflation, ie π_{it}^{e} vs π_{it-1} (β_{1i} vs β_{2i}); the relative importance of the domestic versus the foreign output gap, ie $\tilde{y}_{it}$ vs $\tilde{y}_{it}^{*}$ (β_{3i} vs β_{4i}); the size of the exchange rate pass-through (β_{5i}); and the homogeneity of the β coefficients across countries and country groupings.

Instead of repeating the summary of the results, in the remainder of these comments I will discuss three issues that may have a bearing on the interpretation of the findings.

1. The importance of forward-looking expectations

The terms $\beta_{1i}\,\pi_{it-1} + \beta_{2i}\,\pi_{it}^{e}$ are meant to capture that those who set prices do so in part by looking backwards at the last period's inflation (π_{it-1}) and in part by forecasting how fast prices in general will increase in the future (π_{it}^{e}). If the econometrician finds a high value of β_{2i} relative to β_{1i} it is interpreted to mean that

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forward-looking behaviour is more important than backward-looking behaviour in price setting. This would be a reasonable interpretation if the econometrician could be certain that the variable used to measure π_{it}^{e} captures only forward-looking elements. In the paper, the expected inflation variable, π_{it}^{e}, is measured by consensus forecasts. This is not uncommon in empirical work, but does it necessarily measure only forward-looking elements?

Consensus forecasts are combinations of forecasts of independent professional forecasters. They would presumably construct their forecasts on the basis of how price setters behave and their assessment of the future values of the driving variables of the inflation process.

It follows that professional forecasters will incorporate a backward-looking element into their forecasts in addition to purely forward-looking elements (x_t^e):

$$\pi_{it}^{cf} = \delta_0 + \delta_1 \pi_{t-1} + \delta_2 x_t^e.$$

On this interpretation, the coefficient on the consensus forecast will be a combination of both purely forward-looking elements and backward-looking ones.

2. Measurement of output gaps

In the paper, output gaps are calculated using HP and BN filters. These, as well as other, filters have the property that the gap (at time *t*) calculated with data up to time *t* will in general differ from the gap (at time *t*) calculated with data up to time *T*>*t*. Does this matter for the interpretation of the results?

The answer would be yes if economic agents (forecasters and price setters) use "real time" estimates based on these filters, whereas the econometrician uses the full-sample estimates.

On the other hand, the distinction may not matter if economic agents rely on other indicators in real time that effectively produce gap measures that are consistent with the full-sample estimates the econometrician typically uses ex post.

In general, both measures may be informative in empirical applications, and it would be interesting to try both.

3. Measurement of pass-through coefficients

KMM use nominal effective exchange rates to estimate pass-through coefficients and in general they find that these coefficients are relatively small (albeit not very different from those in Forbes (2019)). A reason for this may be linked to the fact that much of international trade is priced in US dollars (see Boz et al (2017), for example). The relevant exchange rate measure may therefore be the bilateral US dollar exchange rate rather than the effective exchange rate. It would be interesting to investigate whether this is indeed the case.

An additional, possibly minor, issue related to the pass-through estimation is that the exchange rate variable should in principle also include variations in foreign prices in addition to the exchange rate itself: ie $\Delta_4 p_{it}^{*} e_{it}$ and not just $\Delta_4 e_{it}$.

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Robots and labour: implications for inflation dynamics

By Ippei Fujiwara and Feng Zhu[1]

Abstract

This summary investigates how robots affect several variables related to labour, and vice versa. In order to evaluate the causal relationship, we use quality-adjusted robot stock data and labour market data from 26 industries in 33 countries. According to the results obtained from three estimated models – cross-sectional regressions in line with previous studies, panel data regressions and structural panel VAR models – an increase in robot stocks results in higher labour productivity, but has only an ambiguous effect on total employment. Wage increases, but not significantly. Thus, to date, robots should not have exerted a significant influence on inflation dynamics. On the other hand, improvements in labour market conditions lead to significant decline in robot investment. An important lesson obtained in this summary is that when testing whether robots take jobs away from the human workforce, one must also consider the reverse causality from the labour market to robots.

JEL classification: D82, E62, H20, E30.

Keywords: robot, wage, employment, technology.

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1. Introduction

Will machines take jobs away from the human workforce? The impact of "job-stealing" robots has become a growing concern over the last several decades, particularly amid a significant progression in artificial intelligence and algorithms in machine learning. Academics have tackled this issue from a theoretical as well as an empirical angle. Since so-called "robotisation" is a rather recent phenomenon, the data on robots have only been available for the last two to three decades, at most. In addition, compared with the number of theoretical studies on robotisation, only a few empirical studies on the relationship between robots and labour have been conducted.

Graetz and Michaels (2018), Acemoglu and Restrepo (2017a) and Dauth et al (2017) gauge the effects of the increasing usage of robots on the labour market. Implications from these previous studies are somewhat mixed, as shown in Table 1. Graetz and Michaels (2018) evaluate the effect in Europe with cross-sectional data on industries and countries, and conclude that although robotisation increases labour productivity and real wage, it causes no significant effect on labour inputs. Acemoglu and Restrepo (2017a) use US data by commuting zones and conclude that more automation leads to less labour inputs and lower real wage. Dauth et al (2017) focus on Germany, using detailed labour market data, and obtain similar conclusions to Graetz and Michaels (2018), but find no significant effect on real wage at the macro level.

Effect of robots on the labour market

Table 1

	Hours	Employment	Productivity	Wage
Graetz and Michaels (2018)	0	0	+	+
Acemoglu and Restrepo (2017)	–	–	n/a	–
Dauth et al (2017)	0	0	+	0

Notes: +, – and 0 denote positive, negative and insignificant effects, respectively.

This paper, which summarises the findings in our ongoing studies (Fujiwara et al (2019a, 2019c)), also investigates the empirical relationship between robots and labour. In contrast to previous studies, Fujiwara et al (2019a, 2019c) do not have an area focus. There are also substantial differences in data and empirical methodologies.

Regarding the data, there are three primary differences. First, we make use of the quality adjusted robot stock series[2] by using newly available data on the price of robots by industry. Second, we extend the analysis on European countries in Graetz and Michaels (2018) to include China, India, Japan and Korea. Japan and Korea are well known as the frontier countries in robot production and usage. And finally, we use updated EUKLEMS data (significant revisions have been made in the most recent EUKLEMS).

[2] To the best of our knowledge, Fujiwara et al (2019a) and Fujiwara et al (2019) are the first attempts to compute the quality-adjusted robot stock series. For details of quality adjustment, see Fujiwara et al (2019a) and Fujiwara et al (2019).

We also explore new empirical methodologies. Previous empirical studies compare the long-run growth rates of robot stocks with those of labour-related variables, say, over 20 years, with simple cross-sectional regressions. This may lead to significant bias in estimated parameters, if there are country as well as industry-specific fixed effects. Thus, we examine panel data regressions as well. In addition, the causal relationship between robots and labour is investigated by using panel structural vector autoregression (VAR) models.

Our papers, ie Fujiwara et al (2019a, 2019c), have similar conclusions to Graetz and Michaels (2018) and Dauth et al (2017). Advancements in robotics increase labour productivity and wage, but not significantly for the latter. At the same time, only ambiguous effects are observed in labour supply. Thus, we can conclude that robots should not have exerted significant influences on inflation dynamics. Structural panel VAR models show the significantly negative causal effects from improvements in the labour market to robots. This implies the importance in considering the reverse causality from the labour market to robots when testing whether robots take jobs away from the human workforce.

The rest of this summary is organised as follows. Section 2 derives the theoretical relationship between robots and labour using a simple model to obtain empirical implications. In Section 3, our empirical results are summarised. Section 4 concludes.

2. Simple model

This section lays out a simple model to obtain the empirical implications of robots on labour. Consider a social planner's problem to maximise welfare

$$u(C) - v(h),$$

subject to the resource constraint

$$C = f(h, \bar{R}).$$

The variables C, h, and $\bar{R}$ denote consumption, labour and robot stocks, respectively. For the simplicity of analysis, the supply of robots is assumed to be exogenous.

The functional forms are given by

$$u(C) := \log(C), v(h) := \left[(1-\alpha)^{\frac{1}{\varepsilon}} h^{1-\frac{1}{\varepsilon}} + \alpha^{\frac{1}{\varepsilon}} \bar{R}^{1-\frac{1}{\varepsilon}}\right].$$

Without loss of generality, the elasticity of substitution between labour and robots is set to be the same as Frisch elasticity. Then, at equilibrium, we have

$$\mathrm{d}h = (1-\varepsilon) \overbrace{\frac{\alpha \bar{R}^{\varepsilon-2} h}{[(1-\alpha)(1+\varepsilon) h^{\varepsilon-1} + 2\alpha \bar{R}^{\varepsilon-1}]}}^{\oplus} \mathrm{d}\bar{R}.$$

Thus, depending on whether robots and labour are Edgeworth complements ($\varepsilon < 1$) or substitutes ($\varepsilon > 1$), an increase in robots will increase or reduce total employment.

The total differentiation on the relationship between wage, which is equal to the marginal product of labour, and robots leads to

$$\mathrm{d}w = \frac{\overbrace{\frac{1+\varepsilon}{\varepsilon}\left(\frac{h}{1-\alpha}\right)^{\frac{1}{1-\varepsilon}} \alpha \bar{R}^{\varepsilon-2} h}^{\oplus}}{[(1-\alpha)(1+\varepsilon)h^{\varepsilon-1} + 2\alpha \bar{R}^{\varepsilon-1}]} \mathrm{d}\bar{R}.$$

Irrespective of the size of the elasticity of substitution, the effect of robots on wage is always positive.

In the next section, we empirically test whether the above implications obtained in the simple model are observed in the data.

3. Empirical results

For cross-sectional and panel regressions, the estimated equation is given by

$$Y_{c,i,t} = \beta_{1,c} + \beta_{2,i} + \beta_3 Y_{c,i,t-1} + \beta_4 X_{c,i,t-1} + \beta_5 Z_{c,i,t} + u_{c,i,t}.$$

The dependent variables Y are variables in the labour market, where changes in labour productivity is labelled as "productivity"; real wage as "wage"; total employment as "employment" and hours worked as "hours." The explanatory variables X are robot measurements, where changes in the number of robot stock is labelled as "unit"; the percentile measurement in the number of robot stocks as "percentile – unit"; changes in the quality adjusted robot stock as "value"; and the percentile measurement in the number of quality adjusted robot stocks as "percentile – value." Unit and value denote the number of robots and the quality-adjusted robot stocks, respectively. As in Graetz and Michaels (2018), we also use the percentile measurement, which is the percentile of changes in robot units. All observations are classified into 10 groups based on changes in unit measure and are assigned numbers based on the quantile. This is to evaluate the effects at the right tail of the distribution in the changes in robot stocks. The choice of the control variables Z follows Graetz and Michaels (2018). Subscript c, i and t denote country, industry and time, respectively. With cross-sectional regressions, changes are only those between 1999 and 2010 and t as well as fixed effects are not included. Lagged variables are those for the optimally chosen distributed lags in panel regressions.

We also estimate panel bivariate VAR models:

$$\begin{pmatrix} Y_{c,i,t} \\ X_{c,i,t} \end{pmatrix} = B_{1,c} + B_{2,i} + B_3 \begin{pmatrix} Y_{c,i,t-1} \\ X_{c,i,t-1} \end{pmatrix} + B_4 Z_{c,i,t} + C E_{c,i,t}$$

Elements in $E_{c,i,t}$ are orthogonal to each other. C is the lower triangular matrix, which is obtained by the Choleski decomposition of the variance-covariance matrices. The critical assumption in obtaining the structural VAR form as above is the ordering of the endogenous variables in the contemporaneous relationship. We examine all possible combinations of ordering in the bivariate VAR system, namely: (i) when the labour market moves first, ie a shock that increases the labour market variable contemporaneously also affects the robot investment contemporaneously, but a shock that increases the robot investment contemporaneously does not affect the labour market variable contemporaneously; and (2) when the robot moves first, ie a shock that increases the robot investment contemporaneously also affects the labour market variable contemporaneously, but a shock that increases the labour market variable contemporaneously does not affect the robot investment contemporaneously.

3.1. Cross section

Tables 2 and 3 show the estimated coefficients on the robot measurements, namely the effect of robots on the labour market, from cross-sectional regressions with instrumental variables. We employ two instrumental variables employed in Graetz and Michaels (2018). They are "replaceability," ie "an industry-level measure that we call replaceability," and "reaching and handling," ie "a measure of how prevalent the tasks reaching and handling were in each industry, relative to other physical demands, prior to robot adoption."[3] Tables 2 and 3 report the results from "replaceability" and "reaching and handling", respectively. *, **, and *** denote significance at 10%, 5% and 1%.

Cross section with IV: replaceability

Table 2

	Productivity	Wage	Employment	Hour
Unit	–1.289	–1.053	0.298	0.839
Percentile	0.540***	0.501***	–0.029	–0.406***
Value	–1.583	–1.045	0.276	0.667
Percentile	0.629***	0.458***	0.008	–0.301***

Notes: *, **, and *** denote significance at 10%, 5% and 1%, respectively.

With percentile measurements, an increase in robot stocks leads to an increase in labour productivity and real wage. Also, no significant effect is found on total employment. We note that the same result is obtained irrespective of whether the robot stocks are quality adjusted. These all imply that the main conclusion in Graetz and Michaels (2018) is robust even with the updated EUKLEMS data, inclusion of new countries (China, India, Japan and Korea), and quality-adjusted series of robot stocks, as long as the focus is on the right tail of the distribution in the changes in robot stocks.

Cross section with IV: reaching and handling

Table 3

	Productivity	Wage	Employment	Hour
Unit	0	0	0	0
Percentile	0.556***	0.546***	–0.131	–0.393***
Value	4.788	3.885	–1.776	–2.560
Percentile	0.869***	0.530***	–0.125	–0.348***

Notes: *, **, and *** denote significance at 10%, 5% and 1%, respectively.

An increase in robot stocks, however, significantly reduces hours worked. In Graetz and Michaels (2018) and Dauth et al (2017), the effect of robots on labour

[3] We conduct the Cragg and Donald (1993) test for the weak instruments. "Replaceability" turns out to be a weak instrument for all robot measurements, while "reaching and handling" is also a weak instrument for all percentile measurements.

input is close to zero. The negative effect on hours worked is in line with the finding in Acemoglu and Restrepo (2017a).

3.2. Panel

Table 4 summarises the sum of the estimated coefficients over lags on the robot measurements. The estimation period is from 1994 to 2010. Similar to the results from cross-sectional regressions, an increase in robot stocks results in higher labour productivity and real wage, but lower hours. The effect is significant even without resorting to percentile measurements. Contrary to conjectures made in previous studies, a stronger effect is found with annual data than with long-run data. In frequencies higher than 10 to 15 years (but still low), even small changes in robot stocks have a significant effect on the labour market. Altogether, these imply the importance in investigating the impact of robots not only in their right tail and in the long run, but throughout distribution and dynamics.

Dynamic panel — Table 4

	Productivity	Wage	Employment	Hour
Unit	0.005	–0.002	0.005	0.03
Percentile	0.012**	0.006	–0.002	0.03
Value	0.003	–0.008	0.006	0.03
Percentile	0.011*	0.009*	0.002	0.03

Notes: *, **, and *** denote significance at 10%, 5% and 1%, respectively.

3.3. VAR

Table 5 reports the directions of the impulse responses in all VAR models. + and – denote positive and negative reactions, respectively. The sign to the left of the arrow shows the direction at the initial responses, while that to the right of the arrow shows the direction of the responses over the long run. Our primary focus is the sign to the right of the arrow. * denotes the significant responses from zero with one standard deviation.

Directions of impulse responses — Table 5

	Robot shock		Labour shock	
	Robot first	Labour first	Robot first	Labour first
Productivity	+ → +	+* → +	+* → –	– → –
Wage	–* → –*	+* → –	+* → –*	–* → –*
Employment	+* → +	–* → –*	–* → –*	–* → –*
Hours	+* → +*	–* → +	–* → –*	– → –*

Notes: * denotes the significant responses from zero with one standard deviation. The sign to the left of the arrow shows the direction of the initial responses while that to the right of the arrow shows the direction of the responses over the long run.

We can again confirm similar results, in particular, over the long run. Irrespective of the assumptions about the short-run restriction, directions of the long-run effect are the same except for a single case, ie the responses in the total employment to the shock to the robot investment. Graetz and Michaels (2018) and Dauth et al (2017) also report that the responses in the total employment are not significant.

The impact of the robot on the labour productivity is positive, but those on real wage, total employment and hours worked are ambiguous. These are in line with previous studies. The responses of labour productivity are not, however, significant, in contrast to Graetz and Michaels (2018), Dauth et al (2017) and results obtained with cross-sectional and panel data regressions.

On the other hand, regarding robot demand shocks, namely, the direct shocks to the labour market variables, the responses are almost identical irrespective of the identification assumption. The long-run effects are all negative and are significantly negative except for the direct shock to labour productivity, implying a countercyclical robot investment. Improvements in labour market conditions, namely higher wage and labour input, reduce the robot investment significantly. Although we can observe significant effects from labour, the obtained results are not in line with the story offered by the directed technological change of Acemoglu (2002) and Acemoglu and Restrepo (2017b), where the scarcity or tightening of the labour market induces the robot investment. The results from panel VARs hint the importance in considering the reverse causality from the labour market to robots when testing whether robots take jobs from the human workforce.[4]

4. Conclusion

According to the results from three estimated models, while an increase in robot stocks increases labour productivity, it only has an ambiguous effect on total employment. Wage increases but not significantly. Thus, to date, robots should not have exerted significant influences on inflation dynamics. On the other hand, improvements in labour market conditions lead to a significant decline in robot investment, implying a countercyclical robot investment. One important lesson obtained in this paper is is that when testing whether robots take jobs away from the human workforce, one must consider the reverse causality from the labour market to robots.

Several extensions are being considered. First, our results also show large heterogeneity by country and industry in terms of the effects of robotisation. Fujiwara et al (2019b) conduct detailed analysis by industry and country to identify which countries and industries are heavily affected by robotisation. Second, we only estimate bivariate VAR systems, but there are possible endogenous interactions through more labour market variables. Fujiwara et al (2019c) explore the VAR with sign restriction. The directed technological change is tested by identifying a shock that increases wage and hours worked but reduces total employment in the long run and checking the impulse responses to this shock. Third, only the intensive margin

[4] The panel Granger causality test suggests that the developments in the labour market cause robot investment.

adjustments in employment within each sector are considered. Robot manufacturers should increase total employment and wages as robotisation increases. In order to gauge the aggregate effect of robots, we need to consider such extensive margin adjustments through the shifts of labour to the sectors with the most frontier technologies.[5] This topic is covered by Fujiwara, Shirota and Zhu (2019).

[5] We thank our discussant, Yong-Sung Chang for pointing out this important issue.

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Comments on "Robots and labour: implications for inflation dynamics"

Yong Sung Chang[1]

Summary

This paper investigates how the investment in robots affects labour market variables such as productivity, wages and employment, and vice versa. It contributes to this important subject in two ways. First, it enlarges the panel data set that encompass 29 industries and 33 countries by including China, India, Japan and Korea into EUKLEMS. Second, and more importantly, it incorporates the quality of robots by exploiting the price data from *Statistics on manipulator and robots by order, production and shipment* published by the Japanese Robot Association. This price data is applied (ie extrapolated) to other countries assuming that the firms' production technology in different countries is identical as long as they are in the same industry. The paper uses three empirical methodologies: cross-sectional regression, panel regression and panel vector autoregression (VAR). The empirical results establish a strong correlation between the labour market variables and robot investment.

Comments

VAR with sign restrictions

The regression establishes the correlation, but not the causality. The panel VAR, based on the Choleski decomposition, attempts to identify the causality. However, according to the VAR estimates, the results do not significantly depend on the ordering of variables – ie regardless of which moves first, the labour-market or the investment in robots variable. This could mean that the causality runs both ways, or that the ordering restriction based on the contemporaneous correlation does not provide much power in distinguishing the causality. Instead, I would like to suggest an alternative method – the identification of supply and demand shocks based on sign restrictions (eg Canova (2007)).

Firms adopt robots for various reasons. Depending on the nature of production technology, the robots can be a complement or substitute for labour – it can enhance the productivity of labour or replace it altogether. Either way, such technological progress shifts the labour demand curve. Firms may adopt robots because of the high cost of labour (eg a lack of an appropriate workforce or an increase in wage). These events can be interpreted as inward shifts of the labour supply curve in that industry.

Consider a three variable VAR that comprises price of labour (wage or labour productivity), quantity (employment or hours) and robot investment. If the labour

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demand curve shifts, the price (wage) and quantity (hours) would move in the same direction (along the upward-sloping labour supply curve). If the labour supply shifts, the price and quantity would move in the opposite direction (along the downward marginal product of labour). This framework allows us to examine the causes of investment in robots by supply and/or demand factors.

Cross-industry analysis

With the rich data on hand, the authors might be able to test the so-called "routine-biased technological progress hypothesis" put forward by Autor et al (2006) and Acemoglu and Autor (2011). For example, by exploiting the cross-industry variations of employment shares of the routine-task occupation (based on the classification used by Autor et al (2011)), one can ask whether the robot investments have been particularly strong in an industry where a large share of employment engaged in the routine tasks.

Aggregate vs disaggregate analysis

Robots may replace jobs in some industries. However, from an aggregate point of view, the robot-producing firms create jobs. Moreover, history has shown us that industrial revolutions have created new types of jobs that could not be foreseen at the time. Thus, it is important to distinguish between aggregate effect and industry effect. It would be interesting to estimate the aggregate employment effect of robot investment – or the spillover effect of robot investment across industries. Finally, it would be also interesting to estimate the effects of robot investment on the relative wages and/or employment across industries.

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The pass-through from short-horizon to long-horizon inflation expectations

By James Yetman[1]

Abstract

This paper summarises ongoing work that investigates the pass-through from short-horizon and long-horizon inflation forecasts as a way to assess the anchoring of inflation expectations across a sample of 44 economies. It reports an overall decline in the pass-through, with the share of economies having anchored expectations increasing over time. Inflation targeting appears to have played a modest role in improved anchoring. Surprisingly, recent periods with low inflation out-turns are correlated with a decreased pass-through, suggesting that longer-term expectations remain well anchored.

JEL classifications: E31, E58.

Keywords: consensus forecasts, inflation expectations anchoring.

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1. Introduction

Well anchored inflation expectations can play an important role in allowing central banks to pursue an activist monetary policy. One way to assess anchoring is to focus on long-horizon inflation forecasts: more stable forecasts may be indicative of better anchored expectations.

In many economies across the globe, long-horizon inflation forecasts appear to have become more stable over time. Graph 1 displays 6–10 year ahead CPI inflation forecasts collected by Consensus Economics twice a year for 44 economies for as long as they are available. Even for those with a history of high inflation, such as Russia and Turkey, they have tended to fall and stabilise over time.

A complementary way to demonstrate the increased stability of these long-horizon expectations is to compute their standard deviation. Graph 2 displays these for the same data based on five-year (10 observation) rolling samples. These are approximately flat or declining in nearly all cases.

While more stable long-horizon expectations could reflect improved anchoring, there are alternative explanations. For example, the nature of economic shocks could have changed: perhaps they could have become smaller, or less persistent, and this has contributed to a decline in intrinsic uncertainty about future inflation.

One way to disentangle the effects of improving anchoring from other explanations is to focus on how long-horizon expectations respond to specific shocks. Another is to focus on how expectations respond to news. Here we summarise results of ongoing work based on a third approach: the degree to which changes in short-horizon expectations pass through to long-horizon expectations.

This approach makes use of the fact that changes in short-horizon expectations encompass the effects of all types of shocks that influence inflation, including those affecting oil prices, exchange rates and wages. The more similarly long-horizon and short-horizon expectations react, the less well anchored are inflation expectations.

One paper taking a similar approach to ours is Buono and Formai (2018). They estimate an equation in the form $\pi_{it}^{e,l} = \alpha_{it} + \beta_{it}\pi_{it}^{e,s} + \varepsilon_{it}$ for four economies: the euro area, Japan, the United Kingdom and the United States, allowing for time-varying parameters. The left-hand side variable is a longer-horizon forecast, and the right-hand side variable is a short-horizon forecast.

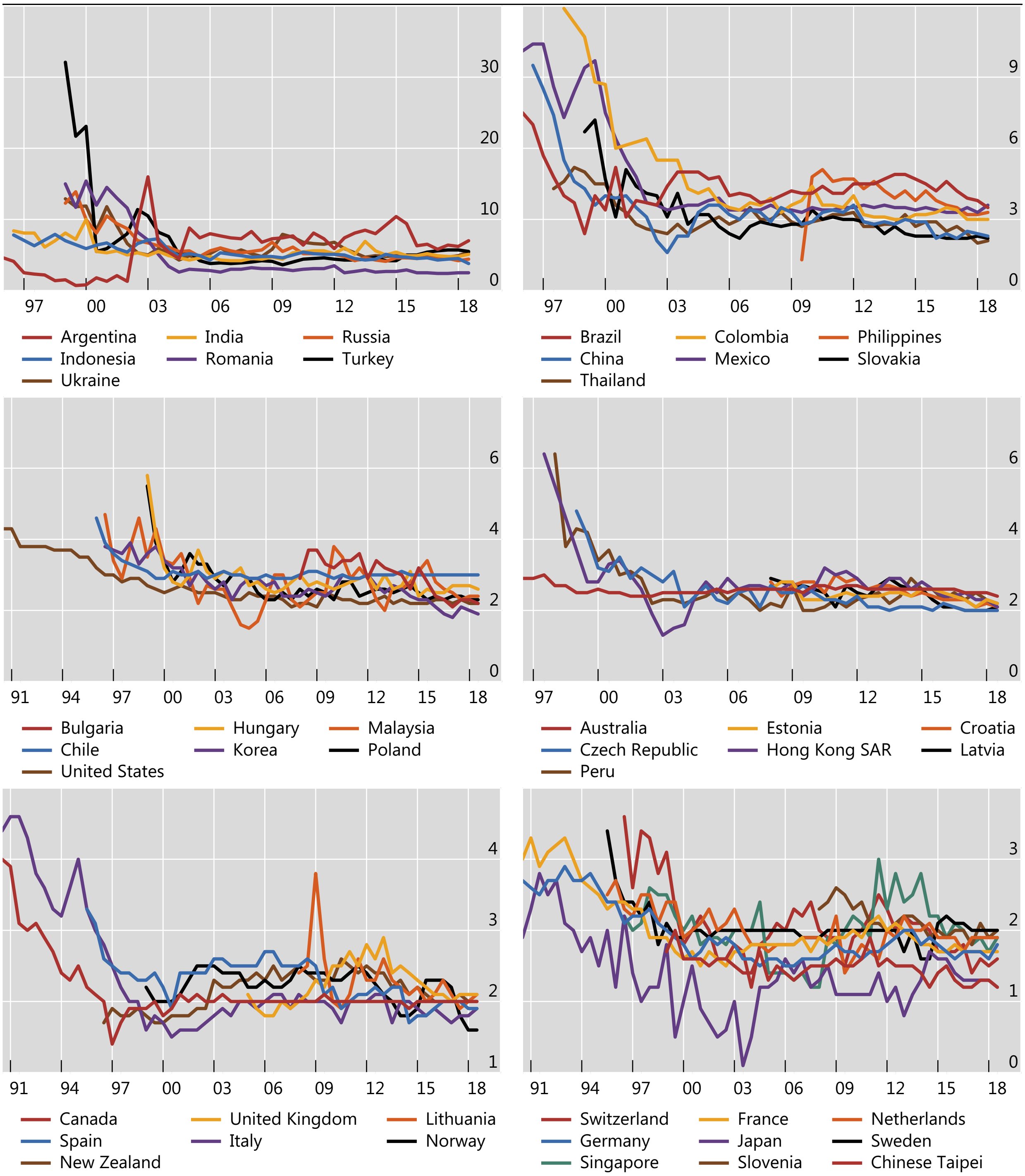
Our approach differs from theirs in several important respects. First, in terms of coverage: we include 44 economies instead of just four.[2] Second, we difference the forecast data, which is arguably necessary in our case since there is evidence of non-stationarity for at least some of our sample. Third, for the long-horizon forecasts we focus on the longest available, 6–10 years ahead, whereas they consider anywhere from two to five years ahead. Fourth, rather than estimating only economy by

[2] For an alternative approach applied to a similarly large set of economies using Consensus Economics forecasts, see Mehrotra and Yetman (2018), who model inflation expectations using a decay function on forecasts with horizons of up to 24 months, where forecasts monotonically diverge from an estimated anchor towards actual inflation as the forecast horizon shortens. They find that this model fits the data well, and indicates that inflation anchors have declined over time for most of the 44 economies in their sample. One limitation of their approach that the current approach addresses is that a 24-month horizon may be too short to assess the anchoring of long-term inflation expectations.

economy, we also consider panel estimation, which allows us to include interactive terms to try to explain what is behind the degree of anchoring and explain its evolution over time.

Long-term inflation forecasts

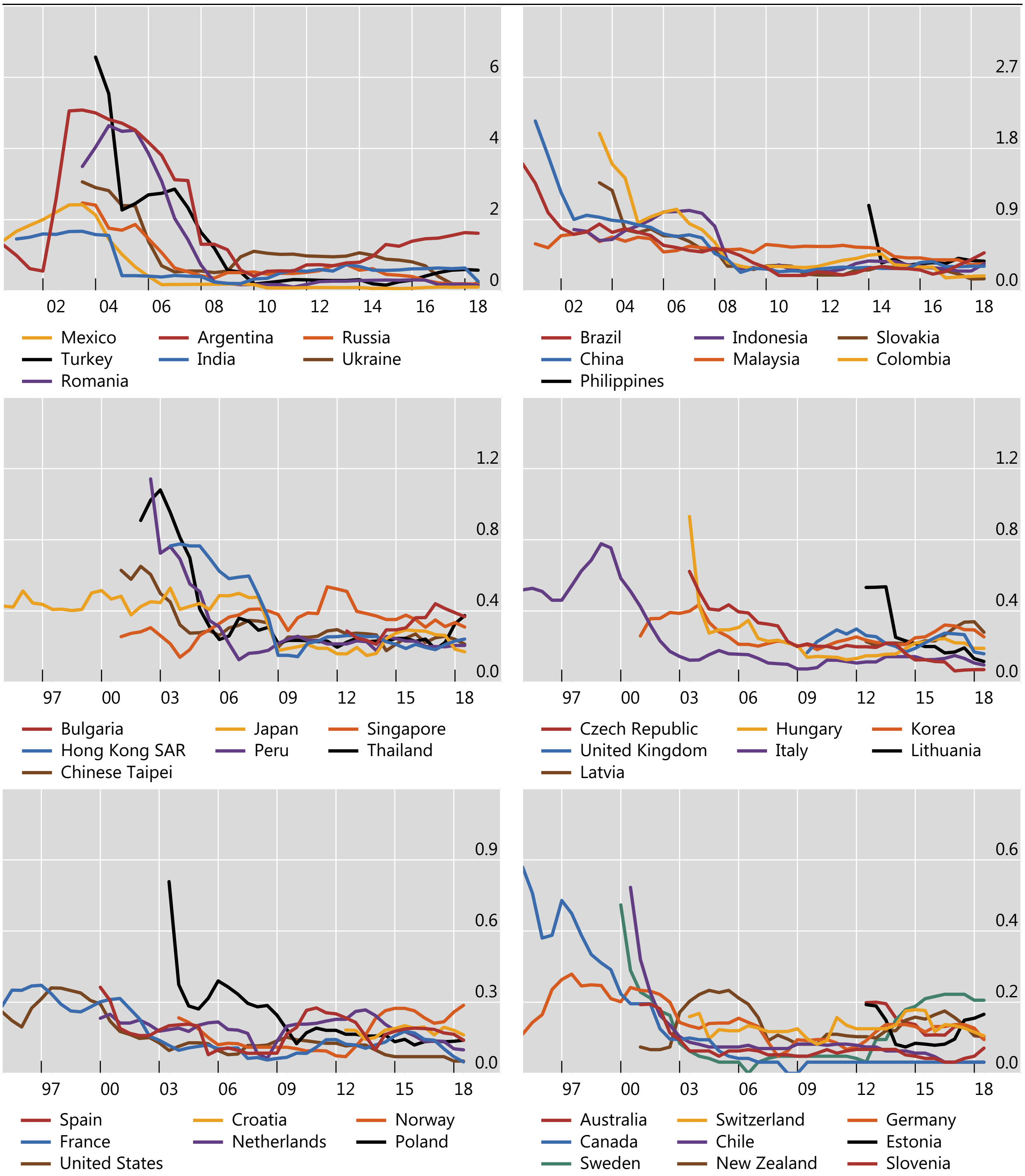
Graph 1



Source: Consensus Economics.

Five-year rolling standard deviation of long-term inflation forecasts

Graph 2



Source: Author's calculations.

We find an overall decline in the pass-through from short-horizon expectations to long-horizon expectations over time in our sample. When we separate the sample into economies that are anchored, contained or unmoored in the spirit of Gefang et al (2012), we note that the share of economies with anchored expectations has steadily improved over the last three decades. We then look to see what might explain this improvement, adding interactive terms to our baseline regression. We find that inflation targeting appears to have played a modest role. However, variables associated with the recent period of low inflation out-turns – low policy rates, persistent deviations of inflation from target and low inflation itself – are correlated with a decline in the expectations pass-through, suggesting that longer-term expectations remain generally well anchored for now.

These questions are especially relevant in the post-Great Financial Crisis (GFC) era, where inflation outcomes have been persistently low in many economies. This work contributes to the existent body of literature investigating the effect of this on anchoring, including Strohsal et al (2016), Miccoli and Neri (2019), Grishchenko et al (2017), Galati et al (2011), Sussman and Zohar (2018), Conflitti and Cristadoro (2018), Garcia and Werner (2018) and Natoli and Sigalotti (2017).

2. Data and estimation

Our data are from Consensus Economics. Each month, Consensus Economics surveys panels of forecasters representing a large number of economies on their forecasts of around 8–10 economic variables for each of the current and next calendar years. In addition, twice per year, they also collect longer-term forecasts of a smaller set of variables for two, three, four, five and 6–10 years ahead. One variable that is nearly always included in these surveys is the percent change in consumer prices (or average annual percent change in the case of 6–10 year forecasts). The availability of these long-horizon inflation forecasts is summarised in Table 1.

For most economies, the shorter-term forecasts are available at the forecaster level, but only averages for the long-term forecasts. We use median short-term forecasts where possible in our study, and the averages published by consensus where not.

Our estimated relationship takes the general form:

$$\Delta\pi_{it}^{e,l} = \beta_i \Delta\pi_{it}^{e,s} + \varepsilon_{it} . \quad (1)$$

The change in the long-term forecast, $\Delta\pi_{it}^{e,l}$, is straightforward to compute. Given that these forecasts are of average inflation 6–10 years ahead of the forecast date, and the forecasts dates are only six months apart, the forecast periods overlap considerably. We simply use the change in the long-term forecasts from one forecast date to the next as our dependent variable.

Long-term inflation forecast availability					Table 1
Economy	Code	Start	End	Notes	AE or EME
Argentina	AR	October 1995	April 2018		EME
Australia	AU	April 1996	April 2018		AE
Brazil	BR	October 1995	April 2018		EME
Bulgaria	BG	September 2007	April 2018	March and September until April 2014	EME
Canada	CA	April 1990	April 2018		AE
Chile	CL	October 1995	April 2018		EME
China	CN	April 1996	April 2018		EME
Colombia	CO	October 1997	April 2018		EME
Croatia	HR	September 2007	April 2018	March and September until April 2014	EME
Czech Republic	CZ	September 1998	April 2018	March and September until April 2014	EME
Estonia	EE	September 2007	April 2018	March and September until April 2014	EME
France	FR	April 1990	April 2018		AE
Germany	DE	April 1990	April 2018		AE
Hong Kong SAR	HK	April 1997	April 2018		AE
Hungary	HU	September 1998	April 2018	March and September until April 2014	EME
India	IN	April 1996	April 2018	Forecasts for fiscal years (end 31/3)	EME
Indonesia	ID	April 1996	April 2018		EME
Italy	IT	April 1990	April 2018		AE
Japan	JP	April 1990	April 2018		AE
Korea	KR	April 1996	April 2018		AE
Latvia	LV	September 2007	April 2018	March and September until April 2014	EME
Lithuania	LT	September 2007	April 2018	March and September until April 2014	EME
Malaysia	MY	April 1996	April 2018		EME
Mexico	MX	October 1995	April 2018		EME
Netherlands	NL	April 1995	April 2018		AE
New Zealand	NZ	April 1996	April 2018		AE
Norway	NO	October 1998	April 2018		AE
Peru	PE	October 1997	April 2018		EME
Philippines	PH	April 2009	April 2018		EME
Poland	PL	September 1998	April 2018	March and September until April 2014	EME
Romania	RO	September 1998	April 2018	March and September until April 2014	EME
Russia	RU	September 1998	April 2018	March and September until April 2014	EME
Singapore	SG	April 1996	April 2018		AE
Slovakia	SK	September 1998	April 2018	March and September until April 2014	EME
Slovenia	SI	September 2007	April 2018	March and September until April 2014	EME
Spain	ES	April 1995	April 2018		AE
Sweden	SE	April 1995	April 2018		AE
Switzerland	CH	October 1998	April 2018		AE
Chinese Taipei	TW	April 1996	April 2018		EME
Thailand	TH	April 1997	April 2018		EME
Turkey	TR	September 1998	April 2018	March and September until April 2014	EME
Ukraine	UA	September 1998	April 2018	March and September until April 2014	EME
United Kingdom	GB	October 2004	April 2018		AE
United States	US	April 1990	April 2018		AE

Notes: unless otherwise stated, long-horizon forecasts are collected every April and October, and are for calendar years. For IN and GB there are also long-term for forecasts for the WPI and RPIX respectively, which we do not use. Forecasts for the euro area are also available beginning in 2003, but we instead focus on the constituent national economies (where available). The final column indicates whether an economy is classified as an advanced economy (AE) or an emerging market economy (EME) in later estimation.

We match these long-term forecasts with the change in the short-term forecasts, $\Delta\pi_{it}^{e,s}$, collected at the same time by Consensus Economics for the same economy, as follows. For each month, there are forecasts for each of the current and next calendar years. Between October and April of the following year, we compute the change in the forecast from October's forecast of the next year to April's forecast of the current year, which are forecasts of the same outcome, but with horizons of 15 months and nine months, respectively, relative to the completion of the year being forecast.

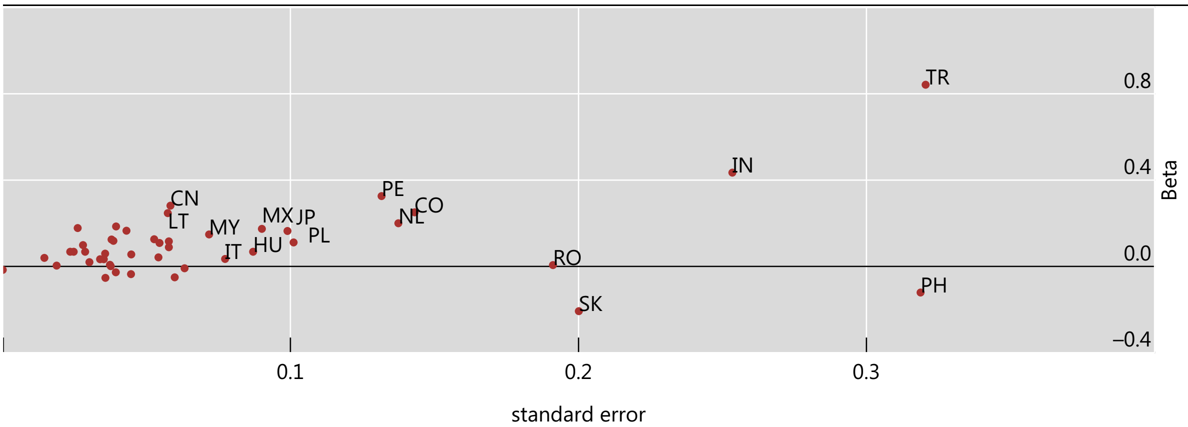
Between April and October, since these are in the same year, and there are forecasts for both the current and next year at each date, we have two possible short-term forecast measures available. The difference between the forecasts of next year's inflation compares horizons of 21 and 15 months, while the difference between the forecasts of this year's inflation compares horizons of nine and three months. We take the average of the two, which has the attractive property of matching, on average, the horizons of the short-term forecasts used to construct the change between October and April.

3. Results

3.1 Evidence of anchoring

We first estimate equation (1) by panel OLS. The model fits well, with an R-squared of 0.88. The coefficients vary widely from low negative numbers to 0.4 for India and 0.8 for Turkey. The standard errors of the estimates also vary widely, with the highest being 0.3 (for the Philippines).

Estimates Graph 3

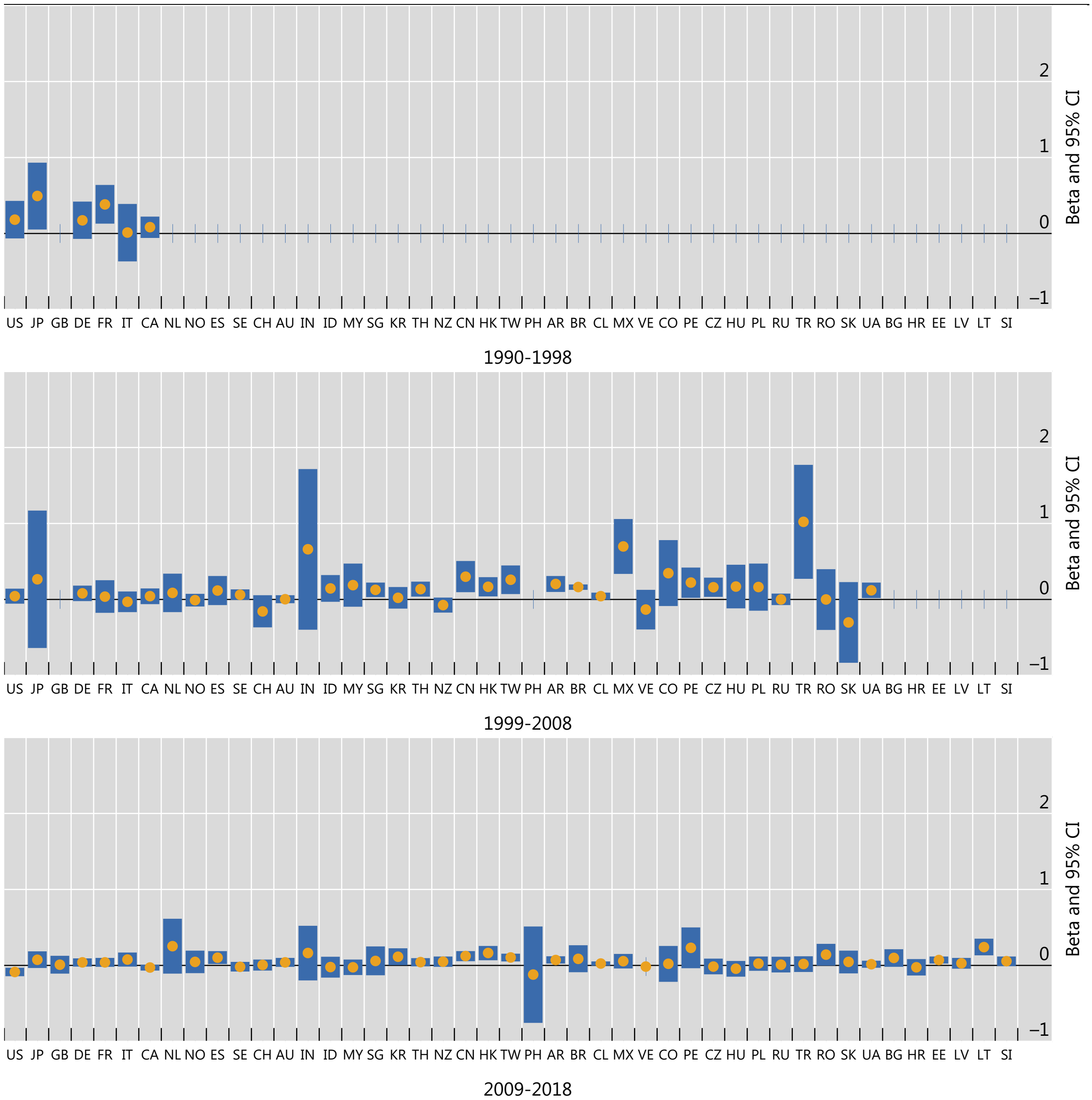


Clearly, well anchored expectations would be expected to result in a low estimate of β_i. But, in addition, we would expect the standard error of the estimate to be small: one can interpret a high standard error as reflecting the fact that there is an uneven relationship between short- and long-term expectations, such that sometimes the pass-through is higher than others. Graph 3 displays a scatter plot of the coefficients and their standard errors for each economy. There is a clear positive relationship between the two.

We also see how the pass-through has changed over time. The results are presented in Graph 4, based on 9–10 year subsamples. We display results for all economies with at least 10 forecast observations (out of a possible 17, 24 or 23 for the respective time-periods). While there are more economies in the sample in later subsamples, these generally display smaller estimates of pass-through (yellow dots) and tighter 95% confidence bands (blue bars).

Expectations pass-through over subsamples

Graph 4



Note: Dots indicate estimated the expectations pass-through for each economy, and bars 95% confidence bands.

We can formalise the improvement in anchoring over time, in the spirit of Gefang et al (2012). We divide economies into one of three categories. Anchored (A) economies are those where the estimated pass-through is low and precisely estimated (which we define as having a 95% confidence interval that includes 0.0 and an upper bound below 0.2). Contained (C) are those that are not anchored but have a pass-through significantly below one (that is, the 95% confidence band excludes 1.0. Finally, unmoored (U) economies are those where the pass-through is not significantly different from one.

Table 2 summarises the results. These clearly indicate an increase in the level of anchoring over time. In the earliest subsample, none of the six economies had anchored expectations. In the middle subsample, 35% of the economies were anchored. If we focus on the six economies that were also in the first subsample and define an improvement in the amount of anchoring as a decline in the pass-through coefficient and a reduction in the standard error, then four of the six show improvement. The other two are mixed (with the pass-through increasing while the standard error decreases, or vice versa). In this middle panel, we have four economies with unmoored expectations: Japan, India, Mexico and Turkey.

Number of economies by degree of anchoring, by subsample — Table 2

	1990–1999	2000–2008	2009–2018
Anchored	0	13	26
Contained	6	20	18
Unmoored	0	4	0
Total	6	37	44

3.2 Understanding inflation expectations pass-through

We next look to see what factors might explain inflation expectations pass-through over time. To do this, we supplement equation (1) above in the following way:

$$/\,\Delta\pi_{it}^{e,l} = \alpha + (\beta_0 + \sum_j \beta_j X_{ijt})\Delta\pi_{it}^{e,s} + \varepsilon_{it} \qquad (2)$$

For X_{ijt}, we consider a number of different potential exponential explanatory variables (j) that could help to explain the degree of the pass-through for short-term expectations to longer-term expectations. These are the average rate of inflation over the past 10 years (MeanInflation) and dummy variables for annual inflation below 1% (LowInflation), a policy rate below 0.3% (LowInterest), an inflation targeting regime (ITDummy, based on the IMF's AREAER), and a stable exchange rate (ERStable, defined as an annual standard deviation against one of the USD or euro (or Deutsche Mark in the pre-euro period) of less than 1% at daily frequency, similar to Carvalho Filho (2010)). In addition, we allow for various combinations of fixed effects.

We display the results in Table 3. The average level of inflation and a stable exchange rate appear to have little effect on the expectations pass-through. Meanwhile, the pass-through falls with low policy rates or low inflation. Inflation targeting also contributes significantly to insulating long-term expectations from changes in short-term expectations in most model specifications. We also note that the intercepts in the specifications without fixed effects are significantly negative, reflecting a trend decline in long-term expectations over time that is not correlated with short-term expectations.

Regression results. All explanatory variables together. Table 3

	All			
MeanInflation	0.000070	–0.00012	–0.000077	–0.00012
	0.26	*0.49*	*0.56*	*0.54*
LowInflation	–0.09	–0.090	–0.10	–0.075
	0.00	*0.09*	*0.00*	*0.10*
LowInterest	–0.11	–0.0060	–0.029	0.12
	0.00	*0.91*	*0.48*	*0.12*
ITDummy	–0.089	–0.10	–0.23	–0.17
	0.01	*0.01*	*0.02*	*0.05*
ERStable	–0.013	–0.014	0.13	0.089
	0.75	*0.74*	*0.01*	*0.30*
Constant	–0.082	–0.067	–0.066	–0.054
	0.00	*0.00*	*0.00*	*0.00*
Year FE	N	Y	N	Y
Economy FE	N	N	Y	Y
N	*1731*	*1731*	*1731*	*1731*
Adj. R2	0.19	0.31	0.30	0.39

NB: p-values in italics; colour coding indicates statistical significance based on robust standard errors, for positive and negative coefficients respectively, as:

0.10,	*0.05,*	*0.01,*	*0.10,*	*0.05,*	*0.01.*

These results indicate that, judging from the pass-through of changes in short-term inflation expectations to long-term inflation expectations, low inflation and low policy rates have not resulted in expectations becoming unanchored. A possible explanation is that forecasters anticipate that these states were/are only transitory, and will be resolved is less time than the forecast horizon of the longer-term forecasts.

We have also examined the anchoring of inflation targeters in isolation, and tested to see if persistent deviations from the target are reflected in a decline in anchoring. Surprisingly, a persistent deviation is associated with a decline in the pass-through that is sometimes statistically significant, consistent with the argument that recent low inflation has not materially impaired anchoring as yet. Meanwhile, having had an inflation target for longer correlates with a decline in the pass-through that is generally statistically significant.

We also compare the results splitting the sample between advanced economies and emerging market economies (EMEs). The one key difference is the inflation targeting dummy: for advanced economies, there is no robust empirical relationship between inflation targeting and the expectations pass-through, while for EMEs inflation targeting regimes experience statistically significantly lower expectations pass-through across all empirical specifications.

4. Discussion and conclusions

In this chapter, we summarise ongoing work investigating the pass-through from short-horizon to long-horizon inflation forecasts as a way to assess the anchoring of inflation expectations. We find an overall decline in the pass-through from short-horizon forecasts to long-horizon forecasts over time in our sample. When we divide our sample into economies with anchored, contained or unmoored expectations, the share of economies with anchored expectations has steadily improved over the last three decades.

We then look to see what might explain this improvement, adding interactive terms to our regression. We find that inflation targeting appears to have played a modest role, especially in EMEs. However, variables associated with the recent period of low inflation out-turns – low policy rates, persistent deviations of inflation from target and low inflation itself – are surprisingly correlated with a decline in the expectations pass-through. This suggests that longer-term expectations remain well anchored: perhaps forecasters perceive low inflation outcomes as transitory, and unlikely to persist for as long as the horizon on the long-term forecasts. However, it remains to be seen if this will continue if low inflation outcomes persist.

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Comments on "The pass-through from short-horizon to long-horizon inflation expectations

by Masazumi Hattori[1]

In this paper, James investigates the evolution of the degree of the pass-through of short-term inflation forecasts to long-term inflation forecasts by using professional forecaster survey data and finds the anchoring of inflation expectations is stronger in many of the sample economies in recent years, especially in ones with inflation targeting in monetary policy. Based on the empirical findings, he judges that the recent missing inflation is transitory.

I think that his explanations on research motivation, data features, estimation methodology and empirical findings are very lucid. The estimations are solid and his interpretation of the estimation results are convincing.

My discussion of his paper consists of four parts. First, I discuss positioning of his paper in the literature. Second, I suggest some possible extensions. Third, I pose some questions. Finally, I point to some caveats that are not specific to his paper but rather general to works in the relevant research fields.

1. Positioning of James's paper in literature

New Keynesian monetary model and pass-through

In the grand literature of New Keynesian monetary theory, a purely theoretical form of the New Keynesian Phillips Curve (NKPC) includes an inflation expectation term and a contemporaneous forcing variable such as unemployment rate and GDP, often in the form of slack, or labour share of income. For empirical estimations, a modified version such as hybrid NKPC that has a lagged actual inflation rate as a persistence term in addition to the two terms in NKPC is used, because of its higher fitting in empirics.

James uses the term pass-through to define the effects of changes in short-term inflation expectations on long-term inflation expectations in a certain economy. Such a concept does not seem to be orthodox in the NKPC; long-term inflation expectations affects current inflation rates, according to the NKPC. My first attempt is to discuss the linkage between the NKPC literature and James's pass-through, thereby explaining the importance of knowing the degree of the pass-through to give an important policy related judgement based on the New Keynesian monetary theory.

The canonical New Keynesian model for empirical studies typically consists of three equations: a hybrid NKPC for inflation rate determination explained above, the dynamic IS curve determining current output (gap) and a Taylor rule describing the

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central bank's decision-making on policy rate, ie short-term nominal interest rate. Mavroeidis et al (2014) calibrated impulse responses to a transitory shock based on the canonical New Keynesian model and show that short-run dynamics of inflation and time necessary for inflation to return to the steady state vary, depending on the value of the parameter for weight on inflation expectations in the hybrid NKPC. A general result is that if the long-run inflation expectations are stable and the weight on inflation expectations in the hybrid NKPC is large, the actual inflation rate will go back to the anchored level in a relatively short time in response to a deflationary transitory shock.

A crucial question is whether the recent actual inflation dynamics, ie missing inflation vis-à-vis shrinking slack in labour market and production potential in many sample economies typically exemplified by the United States, are affecting the long-run inflation expectations; if the answer is "yes", we cannot judge the missing inflation as a transitory phenomenon.

The answer to the question by James's paper is basically "no" or at least "much less than before." The conclusion reinforces the logic of the New Keynesian model for judgment on missing inflation as a transitory phenomenon. This is where James's paper contributes to the literature of New Keynesian monetary theory.

Definitions of anchored inflation expectations

Kumar et al (2015) categorise definitions of anchored inflation expectations. James's interest in this paper is closely related with one of them that is coined "increasingly T-anchored" inflation expectations. This definition means that at a certain time in a time horizon beyond T, every agent's inflation expectations are within a certain small range around the average belief. In plain English, agents' inflation expectations show convergence to the average in the long run from a sampling timing. [2]

A prediction from increasingly T-anchored inflation expectations is "long-term inflation expectations should be unpredictable by using short-term inflation expectations." This prediction is exactly what James investigates.

2. Some possible extensions

In line with the concept of increasingly T-anchored inflation expectations, I would like to suggest a few possible extensions to this paper.

[2] The precise definition of increasingly T-anchored is as follows: Given a sequence $\{\varepsilon_\tau\}_{\tau=0}^{\infty}$ at time t, inflation expectations are increasingly T-anchored at time t if for any horizon $\tau \geq T$, expectations are strongly ε_τ-anchored. Here, "strongly ε-anchored" means that inflation expectations at time t for any horizon $\tau \geq 0$ are strongly ε-anchored if the support of every agent's inflation expectations at that time and horizon lies within ε of the average belief. To be strict, what James's paper is on "quasi" increasingly T-anchored inflation expectations because the survey data, Consensus Forecasts, only include means of professional forecasters' forecasts on long-term inflation rates and the distributions of individual forecasters' forecasts are not available.

More on T in increasingly T-anchored inflation expectations

You can assess the degree of anchoring beyond different forecast horizons. That is, you can choose x-year ahead for T in increasingly T-anchored inflation expectations. Professional forecasters' forecasts on the current year, next year, two-year ahead, three-year ahead, four-year ahead and five-year ahead horizons are available from the same dataset in addition to the six – ten year ahead horizon that James uses. By using two-, three-, four-, and five-year ahead forecasts, you can do more granular analysis, featuring varied categories of sample economies, ie advanced economies vs emerging market economies, economies with inflation targeting vs ones without inflation targeting and so on.

One conjecture is that economies with inflation targeting have a smaller T for the same degree of anchoring than ones without inflation targeting; smaller pass-through coefficients on average for the economies with inflation targeting than the ones without it in a four-year time horizon, for example. As a casual check, you can collect economies with a stable six – ten year ahead forecast levels and compare the trajectories of forecasts between economies with and without inflation targeting; the formers would have a smoother trajectory from the short end to the long end.

More on time variation of the pass-through coefficient for T

For any T, it is possible to use a more advanced methodology than subsample period regressions to investigate the evolution of the pass-through. For example, a time-varying parameter (TVP) regression using a non-linear Kalman filter or a Bayesian estimation is applicable. You can get time series of estimates of the values of the pass-through coefficient for each economy with more precise timings of the change in the values. Then, you can compare the time series of estimates of the coefficients in more detail. For example, one conjecture is that the advanced economies see decreases in the coefficient values earlier than the emerging market economies, reflecting the timings of involvement in the global economy and financial markets or an understanding of the value of price stability.

A data issue arises for this extension. James uses semi-annual survey data. A more precise estimation of time variations in the pass-through coefficient will need higher frequency data. As exemplified in International Monetary Fund (2013), quarterly frequency data will measure up to it. Actually, the quarterly dataset for the same survey from the same survey firm, Consensus Economics, are available for an extra license fee.

3. Some questions

Potential endogeneity between short-term and long-term inflation forecasts

Some factors could affect both short-term and long-term inflation expectations reflected in the forecasts. If this is the case, there could be endogeneity between short-term and long-term inflation forecasts, which could result in bias in the estimate of the pass-through coefficient in the regressions. An appropriate instrument variable to control the potential endogeneity could solve this problem but the regression design in this paper does not resort to the methodology. The potential endogeneity issue is worth considering. This concern is similar to cases of estimations on terms in the NKPC, in which endogeneity between terms is often an issue.

However, my humble opinion is that the issue will not jeopardise James's work. It is highly likely that a missing factor, if any, will affect short-term and long-term inflation expectations in the same direction. I immediately think of the effects of the expansion of shale oil production capacity in the United States that will have deflationary effects on both short-term and long-term inflation expectations. Endogeneity between short-term and long-term inflation forecasts in the same direction in the regressions results in a higher estimated value for the pass-through coefficient, β, than in reality, ie upward bias in the estimate. James showed the estimate of β is small and smaller than before in many economies, thereby concluding the pass-through from the short-term to long-term inflation expectations is to a low or much lower degree there. Hence, the potential upward bias would not counter his conclusion, albeit unprecise estimate if any.

The reason to use short-term inflation forecasts instead of actual inflation rates

I first wondered why James uses short-term inflation forecasts instead of actual inflation rates, or inflation surprises like Kose et al (2019), as the regressor. Is there a superiority issue from the theoretical or empirical perspectives? Or perhaps a technical issue in practice? One supportive opinion for it would be that using the survey data from the same group of respondents will result in consistency in empirical exercises.

In his presentation, James elaborated on the reason that is referred to in the paper; actual inflation rates can be considerably subject to very short-run disturbances while the short-term inflation expectations are expected to reflect fundamentals for current inflation. I now see a more convincing reason.

4. Some caveats

Use of professional forecasters' forecasts

Empirical results in some preceding literature could imply an overestimation of influence of inflation targeting on inflation expectations anchoring in an economy due to the use of professional forecasters' forecasts as variables in the regression analyses in this paper.

Kumar et al (2015) report that firm managers in New Zealand have limited knowledge of the inflation targeting by the Reserve Bank of New Zealand (RBNZ), despite its 25-year tradition; only 12% of respondents to a unique survey knew the correct rate of inflation targeting, ie 2%. Other answers included 1% or 3%" (the bottom and top of the target range) by 25% of the respondents and higher than 5% by 36% of them including 5% respondents choosing 10% or more.

Coibion et al (2018) considered how firm managers revise their macroeconomic expectations in response to new information in New Zealand. The sensitivity of firms' inflation expectations to new information was much higher than for real economic variables. Firms revised their inflation forecasts by most in response to information about the NZRB's inflation targeting, ie the target rate of 2%. Moreover, the effect for the revision dissipated within six months.

Professional forecasters in an economy are the most knowledgeable agents in the private sector on inflation targeting in the economy. If the attraction force of the target rate for agents' inflation forecasts operates, estimations using professional forecasters' forecasts could result in an overestimation of influence of inflation targeting in the economy. If the attraction force of the target rate is stronger for agents' long-term inflation forecasts than their short-run inflation forecasts, which is quite plausible, estimations using professional forecasters' forecasts in light of increasingly T-anchored inflation expectations could result in an overestimation of influence of inflation targeting too.

Transitory… how long?

What is the actual time frame of a "transitory" period? We witnessed missing deflation for about five years immediately following the Global Financial Crisis (GFC) and successive missing inflation for about five years. One convincing background for these phenomena that is consistent with the standard New Keynesian monetary theory is strongly anchored long-term inflation expectations. The question then becomes: What kind of shocks and propagation mechanisms are the background for the transitory phenomena if they are truly transitory? This question is beyond the scope of this paper.

5. Concluding remarks

The paper is crisp and sharp. The paper is an important contribution to the literature on long-term inflation expectation anchoring, providing insight on recent inflation dynamics. Extensions are possible on the time variation of the pass-through coefficient for various forecast horizons. Some general caveats are applicable to this work as well.

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Can an ageing workforce explain low inflation?

By Benoît Mojon (Bank for International Settlements) and Xavier Ragot (Sciences-Po and OFCE)

1. Introduction

We investigate the effects of aging on inflation through the labour market. More specifically, we question whether the weakness of wage inflation over the last decade reflects, at least partially, the increase of labour supply by baby boomers.

The post 2013 recovery of advanced economies has not yet translated into "normal" levels of inflation. Core inflation remains near 1% in the euro area, it has increased from zero to 1% in Japan, while in the United States it is approaching 2% after a sustained recovery. Most other advanced economies also see little inflation. Core CPI inflation, the GDP deflator inflation and wage inflation adjusted for productivity have all remained closer to 1% than to 2%, their pre-crisis nominal anchor. In the case of the euro area, we observe "lowflation" in spite of the creation of over 11 million jobs and over 20 quarters in a row of growth at or above the euro area 1.2 to 1.3 % yoy growth potential.

Unemployment has declined steadily from its peak by several percentage points in the United States, Japan, Germany, Canada, the United Kingdom and Spain. Yet inflation is hardly picking up in these countries. It is very tempting to conclude that the traditional Phillips curve may be broken for good. The weakening of the effects of domestic activity on domestic inflation could result from globalisation or automation. Both weaken the bargaining power of workers. These forces have received a lot of attention.

We instead analyse whether the aging of baby boomers, another well known major transformation of advanced economies, has an impact on inflation. In particular, what has received surprisingly little attention is the tremendous increase in the participation of these baby boomers to the workforce. For instance, 6 of the 7 million jobs created in the euro area between 2013 and 2017 were filled by those aged above 50. In the United States, the share of workers in the workforce aged above 55 has almost doubled from 12% in 1995 to 23% in 2016. In Japan, even the participation of workers aged 65 has increased by nearly 4 million since 2007.

The participation rates of workers aged 55 to 64 has increased from 33% to 55%, on average across OECD countries in the last decade (Graph 1). In Germany, it increased from around 40% until 2003 to above 70% in 2016. This major transformation of the workforce coincides with the setting up of pension reforms as baby boomer cohorts approached the age of retirement. Such demographic conditions may influence the determination of wages drastically.

In principle, this increase in participation may, however, also reflect an increase in labour demand, in which case we expect it would have pushed wages up. The fact of the matter is that, as we show in a technical version of this paper, wages have responded negatively to increased participation of older workers. Therefore, the change in the composition of the workforce is akin to a major labour supply shock by ageing workers. Arguably, these aim to preserve their lifetime purchasing power through postponing their retirement. Ceteris paribus, this positive labour supply shock is likely to push down the levels of wages and unit labour costs. If this transition implies a level shift over several years, it may also impact wage inflation during the years when the transition is taking place.

Our empirical analysis shows that this conjecture in not rejected in the data. Our estimates indicate the participation of the elderly has a specific effect on the labour market. It differs from the one of other age groups. A plausible explanation is that the shorter time horizon of job tenures as retirement approaches reduces the outside value of elderly workers. Hence, they have less incentive to search for other jobs. As a result, the increase of the participation of elderly workers may decrease wage pressure. This conjecture is consistent with a recent analysis of the Bank of Japan (2018), which shows that the wage elasticity of labour supply for the elderly is twice as high as the one of men aged 15 to 64. This in turn contributes to explain why Japanese wages have stagnated in spite of the steady decline of the unemployment rate. This overall negative effect of higher participation by older workers on wage inflation in the period is consistent with the estimates we report in this paper.

2. Data

We assembled annual data on wage inflation, CPI inflation, labour productivity, the rate of unemployment and the participation to labour markets for 19 OECD countries: the United States, Japan, Germany, France, the United Kingdom, Italy, Canada, Australia, Spain, the Netherlands, Belgium, Austria, Finland, Denmark, Norway, Sweden, Switzerland and Portugal and Ireland. Our panel is balanced from 1996 to 2016.

3. Estimation results

The estimates reported in the table correspond to fixed effects panel regressions. Given the narrow cross section of the sample, we also estimated the same equations with the Mean Group Estimators. Results, which are not reported for the sake of space, are very similar to the ones shown here.

As shown in the first column of the table, wage inflation is highly responsive to its three traditional determinants: lagged CPI inflation, productivity and the unemployment rate. The T-stat of the unemployment rate is around 12. The notion that wage inflation is not responding to labour market slack appears extremely unlikely. Of course, the sample we consider here, which includes the Great Financial Crisis, is one with a high degree of co-movement between local labour market conditions and the global business cycle. Hence, some of the effects of the national unemployment rate may imbed the effects of a global slack à la Borio and Filardo (2007). However, while it may be difficult to disentangle the role of the global slack and the local one, Jasova et al (2018) actually show that both domestic and global slack impact domestic inflation. Our estimates show that labour market slack influences wages whichever global or local business cycle is the driver of this labour market slack.

Turning to the period after 2009, for which estimates are reported in the last column, the effects of the unemployment rate on wage inflation are still estimated to be negative. However, given the reduction of the degrees of freedom we have over these seven years, these effects are somewhat less precisely estimated.

The second column reports the panel estimates of a specification augmented with three additional variables:

a. the difference in the unemployment rates of two categories of workers, the ones aged 55 to 64 and the ones aged 25 to 54;

b. the rate of participation to the labour market; and

c. the difference in the participation rates of workers aged 55 to 64 and the ones aged 25 to 54.

Our aim is to assess whether the increased participation of older workers impacts wages. However, we also include in the specification the participation rate of all age groups in order to control for the effects from forces that would drive overall participation rates. Following the same logic, we also assess whether changes in the age composition of unemployed impact wages or the overall effects of the unemployment rate impact wages. Therefore, we include both the overall unemployment and participation rates as well as differences of these age rates above 55 and below 55 to test for a specific effect of the proportion of workers aged above 55 on wage developments.

Among these three additional variables, only the last one has an effect on wage inflation. An increased participation of older workers has a negative effect on wage inflation. We also note that the coefficient of unemployment is hardly affected by including these additional coefficients.

In column four, we include the yearly changes in the difference in participation rates instead of their level. This is to check that our result is not "spurious" given that participation of the elderly shows a trend in many countries while wage inflation rates decline for the sample period. With this specification, we still find a negative effect of participation on wage inflation.

4. Robustness

In Mojon and Ragot (2019), the technical version of this paper, we further investigate the role of labour market developments for the G7 and on wage inflation at the regional level. We put together data on wage inflation, participation of workers aged above 55, CPI inflation and the unemployment rate for 203 European regions. The data cover 17 years from 2000 to 2016.

The panel estimation on the G7 countries is fully consistent with the ones reported here. Henceforth, it is not the case that the negative effects of an ageing workforce on wage inflation is coming from small countries in our OECD panel. It is also observed in the largest among the advanced economies.

Turning to the panel of 203 European regions, we regress the inflation of wage compensation of employees on the participation rate of old workers, lagged inflation and the unemployment rate with regional fixed effects to control for regional heterogeneity. The regressions show that the increase in the participation rate of old workers has a significant and homogeneous negative effect on wage inflation. We also perform regressions controlling for total population, which generates the same outcome. We find that the regional unemployment rate has a negative effect on regional wage inflation, which is a further indication that regional labour markets exibit “Phillips curve-like” patterns.

5. Conclusions

Altogether, the results reported in this paper are reassuring about our understanding of recent labour market dynamics. First, we observe major adjustments of labour supply in response to the ageing of the population. In this respect, the persistent call of central banks for reforms has been either anticipated or answered to by politicians, labour market institutions, employers and workers. This increase in participation implies that potential output should have increased. If the participation of 20% of the working population (population aged 55 to 64 over population aged 20 to 64) has nearly doubled, it means that aggregate output potential should increase as well. It would increase by roughly 6% (as he participation rate of workers aged 55 to 64 increased from 30% to 60%) if the productivity of older workers grows in line with the one of other workers. It should increase by less than 6% if the productivity of older workers is slower. It will, in any case, grow as long as the productivity growth of these older workers is not too negative.

This result also indicates that the Phillips curve transmission of monetary policy to wage inflation is not broken. Increasing participation of older workers has shifted the Phillips curve, which has blurred the response of wage inflation to increasing employment. Central banks who spur activity and employment will eventually harvest domestic wage inflation, and, in all likelihood, inflation of goods and services.

Third, it is not clear yet how high participation rates of older workers will go. We are probably undergoing a very long transition and we don’t know when it will end. But as long as this transition implies a larger slack than measured by the unemployment rate, the economy operates below its “NAIRU” potential. Taking a broader perspective, it seems that the unemployment rate has not been a comprehensive indicator of labour market slack.

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Estimates of wage Phillips curves in a panel of 19 OECD countries

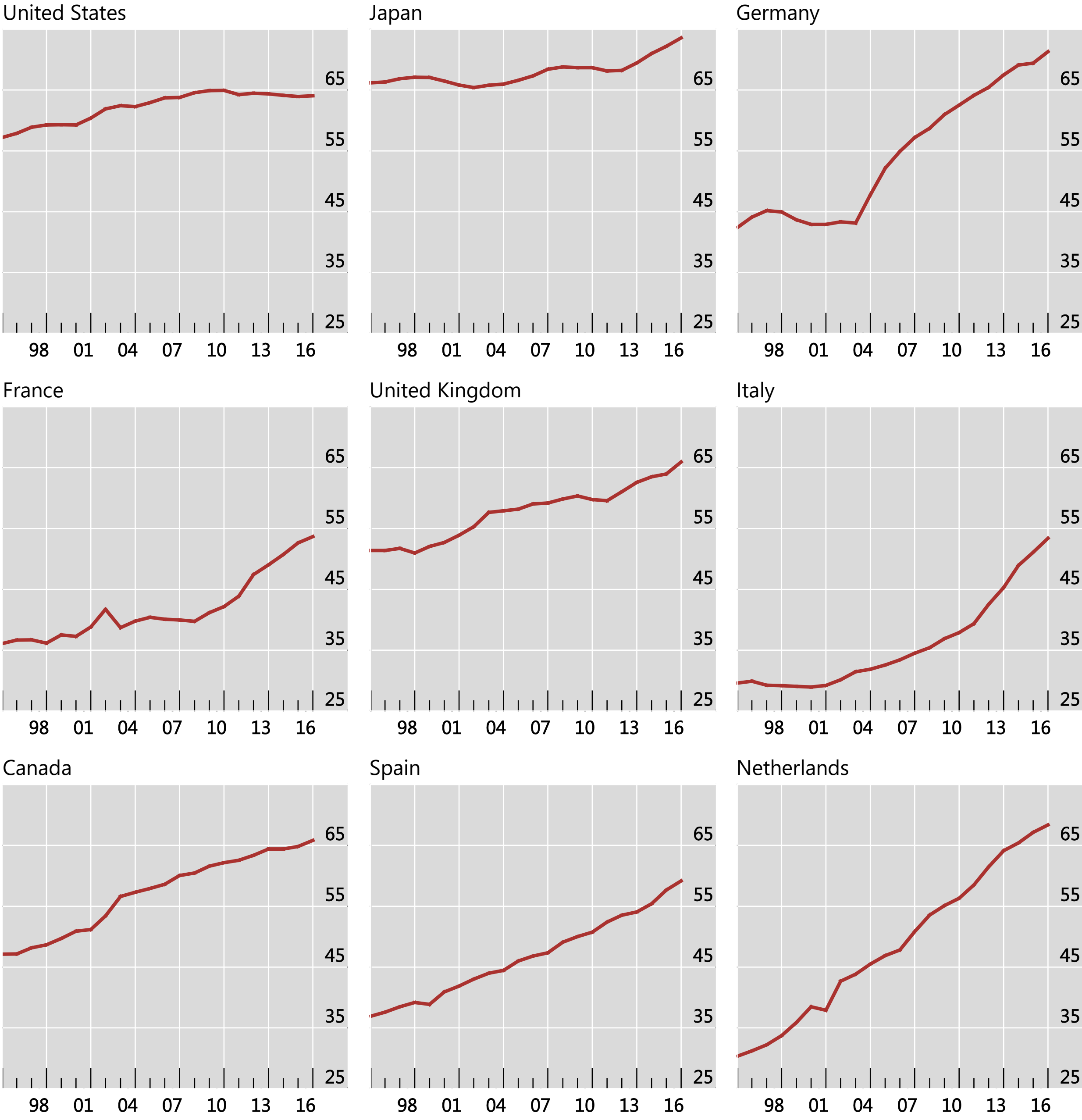
	(1) 1996-2016	(2)	(3)	(4)	(5)	(6) 2010-2016
Lagged CPI inflation	0.417***	0.339***	0.412***	0.340***	0.412***	0,117
	(7.84)	(6.34)	(7.76)	(6.41)	(7.78)	(1.14)
Labour productivity growth	0.229***	0.194***	0.224***	0.194***	0.229***	0.0687
	(6.48)	(5.57)	(6.36)	(5.62)	(6.51)	(1.40)
Unemployment rate	-0.378***	-0.392***	-0.381***	-0.394***	-0.381***	-0.274**
	(-12.99)	(-13.08)	(-12.37)	(-13.96)	(-13.14)	(-3.13)
Unemployment rate difference between the "above 55" and the "below 55" workers		-0.000811 (-0.01)	0.0746 (1.22)			
Participation rate		0.00812	-0.0398			
		(0.26)	(-1.30)			
Participation rate difference between the "above 55" and the "below 55" workers		-0.0666*** (-5.18)		-0.0654*** (-5.63)		-0.0494 (-1.29)
Change in the participation rate difference between the above 55 and the below 55 workers			-0.104 (-1.78)		-0.126* (-2.26)	
Constant	4.289***	2.041	6.919***	2.595***	4.394***	2.392
	(16.46)	(0.94)	(3.51)	(6.62)	(16.68)	(1.95)
Number of countries	19	19	19	19	19	19
Number of observations	425	425	425	425	425	130

The dependent variable is wage inflation. Wages as measured as the compensation of employees.
t-stats are in parentheses.
All regressions include a country fixed effect.
Countries included in the sample: Australia, Austria, Belgium, Canada, Denmark, France, Finland, Germany, Ireland, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom and the United States.

Participation rates: share of “55–64” year old population that work

In per cent of the population

Graph 1



Comments on "Can an ageing workforce explain low inflation?"

Kenichi Sakura[1]

Summary of the paper

This paper investigates the effects of demographic changes on wage inflation. The authors are motivated by three recent phenomena in advanced economies: low inflation, decline of unemployment, and ageing of baby boomers. Specifically, they conjecture that the weakness of wage inflation over the last five years reflects the increase of labour supply by elderly workers (baby boomers). They use a simple model and empirical analysis to show the determinants of wage inflation.

The paper assumes that three types of workers exist in the model: the young, the elderly with continuous careers, and the elderly with discontinuous careers. The average wage in the economy is the weighted average of wages of those three worker groups. On the basis of several stylised facts found in the existing literature, the authors derive the following predictions from a simple model. First, an increase in the participation rate of discontinuous elderly workers decreases the average wage. Second, an increase in the participation of continuous elderly workers has a mixed effect on aggregate wages. Third, wages are more sensitive to the business cycle when the participation rate of discontinuous elderly workers increases and less sensitive when the proportion of continuous elderly workers increases. Overall, the impact of population ageing on the average wage is ambiguous, and the paper attempts to examine it empirically.

In the latter half of this paper, the authors estimate the wage Phillips curve using both country-level and regional-level panel data. The country panel data include 19 OECD countries from 1996 to 2016 and the regional panel data include 203 regions of 24 countries from 1999 to 2016. In addition to the standard wage Phillips curve, the authors include the participation rate of elderly workers (or the difference in the participation rate of workers aged above 55 and those below 55) as an explanatory variable. The results show the negative association between wage inflation and the participation rate of elderly workers, thereby suggesting that the increase in labour supply of elderly workers negatively impacts wage inflation. They also examine this relation by dividing regional samples into two groups according to the pace of population growth. They find it robust even after controlling for possible effects of a shrinking labour force that might cause both lower wages and higher participation rates.

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The views expressed herein are those of the author alone and do not necessarily reflect the official views of the Bank of Japan. The author would like to thank Maiko Koga for her comments and discussions and Enago for the English language review.

Comments

This is an interesting and thought-provoking paper, and the authors discuss a very important topic: the relationship between demographic changes and inflation. Although we observe population ageing and low inflation in many advanced economies, research studies on demographic effects on inflation are relatively limited. The authors contribute to this area of study by empirically showing that an increase in labour supply by baby boomers negatively affects wage inflation rates. From the perspective of making the paper more persuasive, I have three brief comments.

1. Relation of model prediction and empirical analysis

Although the authors present the model to account for the determinants of wage inflation, the connection between the model prediction and the empirical analysis is not well explained.

First, while the model identifies discontinuous and continuous elderly workers, the empirical analysis does not consider this distinction. To fill this gap, it might be more persuasive to add an empirical exercise using employment data by job tenure (duration of employment).

Second, although the authors have derived a prediction regarding the business cycle effects on wages, they do not test it in their empirical section. According to their model, the "slope" of the wage Phillips curve would also be affected by the labour supply of elderly workers; for example, wages are more sensitive to the business cycle when the participation rate of discontinuous elderly workers increases. It would be interesting to verify this prediction, for example, by introducing an interaction term of the participation rate of elderly workers and a variable representing economic slack.

2. Specification of wage Phillips curve

The paper presents the wage Phillips curve in which wage inflation is determined by lagged inflation, labour productivity growth, unemployment rate and participation rate. One concern is that it is unclear what each term represents. It would be more persuasive if the authors could explain on which model their specification is based.

For example, in a recent work on the wage Phillips curve, Gali (2011) gives a similar specification based on the standard New Keynesian model. In his model, wage inflation is associated with price inflation, unemployment and productivity growth. He theoretically shows that wage inflation is proportional to the discounted sum of expected deviations of current and future average wage markups from their desired levels. Assuming wage indexation to price inflation and showing the relationship between the wage markups and the unemployment rate, he explains that price inflation, unemployment and productivity growth are relevant variables.

3. Alternative explanation

Some previous studies analyse demographic effects on inflation on the basis of other mechanisms. Therefore, it should be noted that the observed correlation between population ageing and low inflation could also reflect other factors than those the authors point out.

One alternative explanation is based on the secular stagnation hypothesis (Summers (2013)). Eggertsson et al (2019) formally model a mechanism in which a decline of population growth decreases loan demand, thereby resulting in downward pressure on the natural interest rate. As the lower bound of nominal interest rates becomes binding, the impact of monetary policy declines and this could lead to longer recessions and lower inflation.

Second, Katagiri et al (2019) describe the association between demography and inflation from the political economy perspective. In their model, population ageing stemming from an increase in longevity leads to deflation by increasing the political influence of the older generation. Under the fiscal theory of the price level (FTPL) setting, the government increases its income tax rates (which harms the young) to avoid an increase in prices (which would harm the elderly).

Third, Fujita and Fujiwara (2016) develop a New Keynesian search/matching model to demonstrate that changes in demographic structure induce skill (productivity) heterogeneity in the workforce, which results in low-frequency movements in the real interest rate. When monetary policy follows the standard Taylor rule that fails to internalise the time-varying nature of the natural interest rate, the economy experiences low-frequency movements in the inflation rate as well.

Finally, other studies have also attempted to explore empirically the relation between age structure in the economy and inflation (eg Juselius and Takáts (2015); Yoon et al (2018); Barbiellini Amidei et al (2019)).

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Strategic complementarity and asymmetric price setting among firms

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Abstract

Using a large panel of firm survey data from Japan (*Tankan* survey), we provide micro evidence of strategic complementarity in firms' price setting. We find that a firm's price adjustment is affected by its competitors' pricing behaviour and that this adjustment is larger when the firm is lowering its price, which accords with the theoretical predictions of quasi-kinked demand. Our results also indicate that firms with greater pricing power tend to be less sensitive to their competitors' behaviour. Finally, we observe that heightened demand uncertainty mitigates the effect of shifts in demand conditions on the likelihood of price adjustment.

JEL classification: D22, D84, E31, E32.

Keywords: demand uncertainty, firm survey data, price setting, strategic complementarity.

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1. Introduction

Moderate inflation has been a pervasive phenomenon in advanced economies in the past two decades (see, for example, Blanchard et al (2015), IMF (2013)). Japan is often cited as an illustrative case, having struggled with prolonged deflation for a decade and a half. With this in mind, we investigate firms' price-setting behaviour using firm survey data from Japan. Our conjecture is that prolonged deflation may be attributed to firms' asymmetric price setting: firms may refrain from increasing their own prices because their competitors are doing the same. These interactions in pricing attitudes among firms can be described as price setting under a quasi-kinked demand curve. In this setting, theory predicts: (1) a price increase (decrease) by competitors makes it optimal for a firm to increase (decrease) its own price, so that firms' pricing decisions are mutually reinforcing; and (2) firms' reactions to their competitors' prices are asymmetric: they tend to be more responsive to price reductions by competitors than to price rises, and as a consequence, firms are more cautious about decisions to increase prices compared to lowering them.[3] The aim of our paper is to provide micro evidence to support such asymmetric price setting as predicted by the theory.[4]

Data and main findings

Our firm survey data are from the "Short-term economic survey of enterprises in Japan" run by the Bank of Japan and known as the "*Tankan* survey". This covers approximately 10,000 firms in Japan and boasts excellent quality sampling, achieving a response rate of almost 99% from firms across a wide range of sectors. Results from analysing the data can thus be a reliable source of inference regarding the macroeconomy.

The advantage of the survey is that it allows us to identify competing firms for a given firm, using their reporting about the main products and services. We classify firms into 636 industry categories, which is broadly consistent with the 4-digit industry level in the Japan Standard Industrial Classification.

Using the data, we find the following results. First, we find that firms' price setting responds to their competitors' prices – evidence of strategic complementarity. Although the assumption of strategic complementarity in pricing is standard in the New Keynesian literature (Woodford (2011)), the empirical evidence to support this is scarce. Our study provides micro evidence in this respect. Second, the degree of strategic complementarity is stronger for price decreases. This asymmetry in the reaction to competitors' prices is consistent with the theoretical predictions for firms' price setting under a quasi-kinked demand curve. Third, we find that when firms hold

3 Strategic complementarity also directly affects the slope of the Phillips curve describing the relation between price changes and output; specifically, it acts to weaken this relation. In other words, the greater the degree of strategic complementarity, the less the price responds to an unexpected variation in nominal spending (Woodford (2011)).

4 Our analysis is also motivated by an attempt to identify particular sources of real rigidities. Studies on real rigidities describe persistent real effects being generated from nominal shocks. Levin et al (2008) demonstrate the importance of exploring the mechanisms underpinning these effects, since different mechanisms may lead to different implications for monetary policy even under equivalent New Keynesian Phillips curves. They also argue that utilising micro data reveals insights into the economic structure and implications that could not have been obtained from macroeconomic data alone. We exploit a large set of firm panel data to directly examine the presence and the source of real rigidities at the firm level – that is, the existence of strategic complementarity in pricing.

higher inflation expectations, they are more likely to raise the prices of their own goods. Fourth, by extending our main analysis, we also observe that firms are heterogeneous in their degree of strategic complementarity depending on their market share in an industry. Firms with higher pricing power tend to exhibit less caution in responding to competitors' price changes. Fifth, we find that demand uncertainty also affects pricing. Heightened demand uncertainty mitigates the demand effect on price adjustment probabilities, a phenomenon which we take to be evidence of "wait and see" pricing.

Related literature

Our study is related to three strands of literature. The first of these is the literature on firms' price-setting behaviour, where current wisdom basically favours state-dependent pricing over time-dependent pricing (eg Klenow and Kryvtsov (2008), Nakamura and Steinsson (2008), Klenow and Malin (2010), Honoré et al (2012)). These studies demonstrate that price setting depends on *aggregate* variables. For example, Honoré et al (2012) find that a rise in inflation encourages firms to increase prices.[5] However, there are few studies exploring whether pricing decisions are dependent on *firm-specific states*. Notable exceptions are Lein (2010) and Amiti et al (2019). Lein (2010) presents evidence from Swiss manufacturing firms that pricing decisions rely on firms' current situations including the cost of intermediate goods. Amiti et al (2019) find that firms react to competitors' price settings using data on Belgian manufacturing firms. Our contribution to the literature is to provide new evidence of the asymmetric reaction of firms to competitors' prices, consistent with expected behaviour under a quasi-kinked demand curve, and to help explain why price increases occur less often than price decreases. This evidence is confirmed across a broad range of industry categories including non-manufacturing firms.

Second, our study is associated with the literature on the kinked demand curve and, more generally, work on variable elasticity of demand. Though constant elasticity of demand is still the most popular setting in macroeconomic models, a growing literature demonstrates that kinked demand, among other forms of variable demand elasticity, is a useful theoretical framework to account for real rigidities in which nominal shocks generate persistent real effects (eg Kimball (1995), Klenow and Willis (2016), Shirota (2015), Kurozumi and Van Zandweghe (2018)). However, micro evidence to support this setting is still limited. A notable exception is Dossche et al (2010), who provide such support using supermarket scanner data for price and quantity at the *goods level*. Our paper contributes to the literature by offering empirical support for the validity of the kinked demand curve theory at the *firm level*.

Third, our study also builds on a growing literature that uses firm survey data to demonstrate how heterogeneous expectations among firms result in diversified behaviour. Bachmann et al (2013) demonstrate that firm-level uncertainty leads to a significant reduction in production using firm survey data from Germany. Using Japanese firm-level data, Koga and Kato (2017) reveal how heterogeneity in firms' expectations regarding industry demand growth affects investment decisions. Tanaka et al (2019) study the relationship between the accuracy of macroeconomic forecasts and firm performance. Morikawa (2016) also demonstrates the negative relationship between subjective uncertainty and investment. In a similar vein, we explore how firms' heterogeneous expectations are reflected in their own pricing decisions.

[5] The authors provide empirical support for the prediction of Ball and Mankiw (1994).

2. Description of the survey data

The data we use are from the "Short-term economic survey of enterprises in Japan" (widely known as the *Tankan* survey) conducted by the Bank of Japan. The survey aims to provide an accurate picture of business trends among firms in Japan to support the appropriate implementation of monetary policy. The survey is conducted quarterly, in March, June, September and December, across broad industry categories. The survey population comprises private firms excluding financial institutions in Japan with a capital of JPY 20 million or more, and totals approximately 220,000 such firms. Sample firms are selected from the survey population based on industry and size classifications so as to meet the criteria for statistical accuracy, and the number of sample firms is about 10,000.[6] Firms are classified into 31 industry groups and three size groups. Industry groups are based on the Japan Standard Industrial Classification released by the Ministry of Internal Affairs and Communications. Size groups are based on capital size reported by firms: large enterprises (with capital of JPY 1 billion or more), medium-sized enterprises (with more than JPY 100 million but less than JPY 1 billion in capital), and small enterprises (with more than JPY 20 million in capital, but less than JPY 100 million).

Firms report assessments of 10 items including business conditions, employment conditions, demand conditions in each industry, changes in output prices, and changes in input prices. Answers are provided for two horizons: for the current quarter and the next quarter. They also report annual projections for sales, profits, fixed investment and so forth. The items on which we focus in this paper are firms' assessments of pricing attitude and related factors. In the question on pricing attitude, firms are required to choose from three possible responses: rise, unchanged and fall, and allocated scores for each response are 1, 0 and –1, respectively.

In addition, we adopt items that we presume reflect real marginal cost: namely, firms' views of changes in input price $(Cost_{i,t})$ and changes in employment conditions $(LC_{i,t})$. We use employment conditions to represent labour costs on the assumption that firms answering "excessive employment" tend to face lower labour costs. $Demand_{i,t}$ constitutes the demand conditions reported by firms in each industry.

Competitors' prices for each firm $ComPrice_{i,t-1}$ are calculated from the survey responses by computing the average score of the pricing attitudes of $N-1$ other firms in the same industry (excluding firm j itself). We then take this figure to capture pricing attitudes of firm $j's$ competitors: that is, $dp_{-i,t} = \sum_{j \neq i,t} dp_{j,t}/(N-1)$. We then use the one-period lag of competitors' prices $ComPrice_{i,t-1}$ in our estimation. To identify the competitors for a given firm, we produce 636 specific industry categories based on firms' reporting about their main business products and services. This classification is broadly consistent with the four-digit industry level in the Japan Standard Industrial Classification.[7] When we compute the variable for competitors' prices, we use a one-period lag of competitors' prices $ComPrice_{i,t-1}$ as the explanatory variable for desired price at time t to mitigate the endogeneity issue, assuming that firms react to competitors' prices one quarter later. As we cannot

[6] As a large number of sample firms are added in regular revisions conducted at three- to five-year intervals, our original data set is an unbalanced panel covering approximately 14,000 firms. The number of firms we use in the analysis is trimmed down to about 10,000 by computing the variable for competitors' prices, as explained later.

[7] The average number of competitors per industry is 12.78.

define competitors' prices for firms with no competitor in their market, such firms are excluded from the estimation.[8]

3. Empirical specification

Dotsey and King (2005) model the price-setting behaviour of firms under the quasi-kinked demand curve theory proposed by Kimball (1995).[9] The theoretical predictions derived from the model are as follows: (1) each firm's desired price depends on firm-specific states including its competitors' prices, real marginal costs, and demand; (2) the desired price is also affected by the aggregate inflation rate; and (3) the effect of competitors' prices is stronger for price reductions than for price increases. The aim of this paper is to examine the above predictions using a large panel of firm survey data.

To test the above theoretical predictions, we use survey responses to capture firms' pricing attitude. We model the relationship between $P^*_{i,t}$ and $PriceChange_{i,t}$ as in a limited dependent variable model, where $P^*_{i,t}$ is a latent variable describing changing prices, and we only observe $PriceChange_{i,t}$, which is a qualitative response on firms' pricing stances. When $P^*_{i,t}$ exceeds the threshold θ_2, firms report a price increase; when $P^*_{i,t}$ drops below θ_1, they report a decrease; in all other cases, they report that prices are unchanged.[10]

$$Price_{i,t} = \begin{cases} 1 \text{ if } \theta_2 \leq P^*_{i,t} \ (Increase) \\ -1 \text{ if } P^*_{i,t} < \theta_1 \ (Decrease) \\ 0 \text{ if } \theta_1 \leq P^*_{i,t} < \theta_2 \ (Unchanged) \end{cases}$$

We assume that the desired price is determined in the following way:

$$P^*_{i,t} = \beta_1 Cost_{i,t} + \beta_2 LC_{i,t} + \beta_3 Demand_{i,t} + \beta_4 ComPrice_{i,t-1} + \gamma_1 x_i + \gamma_2 y_t + \epsilon_{i,t},$$

where $P^*_{i,t}$ is the desired price. $Cost_{i,t}$ and $LC_{i,t}$ are variables representing the cost of intermediate goods and the cost of labour, respectively. We regard these variables as reflecting real marginal cost. $Demand_{i,t}$ describes the demand conditions reported by firms in each industry. All of the above variables are obtained from the survey responses, and take one of three possible values: 1, 0 or –1.

The variable for competitors' prices, $ComPrice_{i,t-1}$, is calculated from the survey responses as explained in the previous section.[11] Amiti et al (2019) address the

[8] The number of sample firms in this paper is trimmed down to about 10,000 firms. This sample covers 72.7% of the original survey data in terms of the number of firms and 95.7% on a sales basis as of March 2015.

[9] Koga et al (2019) derive the linear relations among the variables, since Dotsey and King (2005) do not explicitly show how each firm's desired price is affected by relevant factors in a linearised form.

[10] We describe response categories as "Increase", "Decrease" and "Unchanged", while the original categories are "Rise", "Fall" and "Unchanged".

[11] The use of the one-period lag relies on the assumption that firms react to competitors' prices one quarter later. This specification does not fully measure the effect of competitors' prices, since we do not capture faster or slower reactions to others' price changes. Thus, the effect that we estimate for competitors' prices may be considered a lower bound for observable relations among competing firms.

endogeneity issue by estimating the model with instrumental variables. They use imported components of the firm's cost as an instrumental variable for competitors' prices. In our analysis, as we do not have a suitable candidate for instrumental variables, we use the one-period lag of competitors' prices. Inflation expectations $\left(EInflation_{i,t}^{l_{yys}}\right)$ affect the desired price in our setting. However, as inflation data are only available from 2014, we omit this variable in the benchmark case of our estimation and add it later with a more restricted sample. x_i is the vector of firm-specific variables. As there may be industry-specific or size-specific factors that affect not only firms' pricing attitudes but also cost and demand factors, x_i includes industry dummies for the 31 industry categories and size dummies for the three capital size categories.

y_t includes fixed effects for the year and two additional dummies to control for the effects of specific aggregate shocks such as the Great Financial Crisis (Q3 2008–Q1 2009) and the consumption tax increase (Q2 2014). The estimation is conducted for the period from Q1 2004 to Q4 2017. As a large sample revision was carried out in Q1 2004, the quality of the estimation might be impaired if we used older historical data. In robustness checks, we also consider whether there is a time-dependent pricing factor by adding Taylor dummies. These are dummy variables that denote when the last price change occurred, between one and eight quarters ago. Quarter dummies are also added to control for price changes occurring in a specific quarter. Our main results are robust to these controls for time dependency, as shown in Appendix Tables 1 and 2 of Koga et al (2019).[12]

4. Main results

Baseline

This section describes the main results. All tables report average marginal effects. Standard errors are clustered at the firm level. Table 1 shows the regression results for price changes. The dependent variable is $PriceChange_{i,t}$ and the estimation is conducted by the ordered probit model.[13]

Variations in columns give the results for alternative measures of marginal costs. Columns (1) and (2) show the results using firms' responses for current input prices and forecast input prices, respectively; column (3) shows the results when using both. Column (4) shows the results of the correlated random effect model that controls for firm-specific heterogeneity, using historical averages of independent variables for each firm.[14] The results listed in column (1) are regarded as the benchmark case hereafter.[15]

[12] Appendix Tables 1 and 2 are not reported here, and are available in the working paper version of our paper (Koga et al (2019)).

[13] The dependent variable, $PriceChange_{i,t}$, takes the values of 1 (increase), 0 (unchanged) and –1 (decrease).

[14] The correlated random effect model works as a robustness check when firm fixed effects cannot be controlled for in the probit model. A detailed explanation is provided in Wooldridge (2011).

[15] When we use both current and forecast input price choices simultaneously, the coefficient on the latter is negative. This may reflect the correlation between current and forecast input prices. A similar result is found in Lein (2010).

The coefficient on current input price is significantly positive in all cases. The coefficient on employment conditions is significantly negative, suggesting that higher labour costs also push up the probability of a price change. The demand factor also contributes to the likelihood of price changes. Even after controlling for these factors, however, competitors' prices are positively correlated with the firm's pricing stance. Coefficients on all of these variables are statistically significant.

As for the magnitude of the impact, the estimates suggest that when average scores of input prices and competitors' prices show a one-unit increase, the probability of an adjustment in output prices is about 8 and 10 percentage points higher, respectively.

Baseline results Table 1

		Price changes			
Sample period: Q1 2004～Q4 2017		(1)	(2)	(3)	(4)
Cost_t	(+)	0.080***		0.084***	0.081***
		(0.002)		(0.002)	(0.002)
LC_t	(−)	–0.015***		–0.009***	–0.015***
		(0.001)		(0.001)	(0.001)
$\mathrm{E}_t(\mathrm{Cost}_{t+1})$	(+)		0.038***	–0.008***	
			(0.001)	(0.001)	
$\mathrm{E}_t(\mathrm{LC}_{t+1})$	(+)		–0.015***	–0.009***	
			(0.001)	(0.001)	
Demand_t		0.042***	0.046***	0.042***	0.041***
		(0.001)	(0.001)	(0.001)	(0.001)
$\mathrm{ComPrice}_{t-1}$	(+)	0.099***	0.117***	0.099***	0.100***
		(0.003)	(0.003)	(0.003)	(0.003)
Dummy: Great Financial Crisis		–0.005***	–0.006***	–0.005***	–0.005***
		(0.001)	(0.001)	(0.001)	(0.001)
Dummy: Consumption tax		0.001	0.003**	0.001	0.001
		(0.001)	(0.002)	(0.001)	(0.001)
Adjusted pseudo R-squared		0.136	0.117	0.136	0.133
Observations		409,736	409,736	409,736	409,736
Number of IDs		10,300	10,300	10,300	10,300
Size fixed effect		Yes	Yes	Yes	Yes
Industry fixed effect		Yes	Yes	Yes	Yes
Year fixed effect		Yes	Yes	Yes	Yes
Correlated random effect		No	No	No	Yes

Note: Standard errors in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

Asymmetry

Table 2 shows the results for asymmetric price setting, based on subsamples of firms reporting, respectively, price increases and price decreases, and the estimation is conducted utilising the probit model.[16] In this specification, the expected signs are reversed in the two halves of the table. For example, an increase in input price pushes up the probability that a firm raises its price, and pushes down the probability of a price reduction. According to our theoretical setup in the previous section, a firm is expected to react asymmetrically to competitors' prices. The empirical results bear out this prediction. The most substantial difference is observed in the average marginal effects of competitors' prices, and the effect of a price decrease is more than double in absolute value the effect of a price increase. These results support the premises of price setting under a quasi-kinked demand curve.[17]

Results: asymmetry

Table 2

1: increase (decrease) 0: otherwise		Price increases		Price decreases	
Q1 2004～Q4 2017		(1)	(2)	(3)	(4)
Cost_t	(+)	0.095***	0.095***	–0.070***	–0.072***
		(0.002)	(0.002)	(0.002)	(0.003)
LC_t	(−)	–0.009***	–0.009***	0.030***	0.030***
		(0.001)	(0.001)	(0.002)	(0.002)
Demand_t		0.020***	0.021***	–0.085***	–0.083***
		(0.001)	(0.001)	(0.002)	(0.002)
$\mathrm{ComPrice}_{t-1}$	(+)	0.067***	0.067***	–0.149***	–0.150***
		(0.002)	(0.002)	(0.005)	(0.005)
Dummy: Great Financial Crisis		0.008***	0.008***	0.013***	0.013***
		(0.002)	(0.002)	(0.003)	(0.003)
Dummy: Consumption tax		0.001	0.001	0.006	0.006
		(0.002)	(0.002)	(0.004)	(0.004)
Adjusted pseudo R-squared		0.152	0.152	0.184	0.224
Observations		409,736	409,736	409,736	409,736
Number of IDs		10,300	10,300	10,300	10,300
Size fixed effect		Yes	Yes	Yes	Yes
Industry fixed effect		Yes	Yes	Yes	Yes
Year fixed effect		Yes	Yes	Yes	Yes
Correlated random effect		No	Yes	No	Yes

Note: Standard errors in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

Empirically, a similar asymmetry shows up also for labour costs and demand. The absolute values of the coefficients on $LC_{i,t}$ and $Demand_{i,t}$ are larger for price decreases than for price increases. This is broadly consistent with the implications of the quasi-kinked demand curve: when firms face higher labour costs or demand and

16 In the price increase (decrease) case, the dependent variable $PriceIncrease_t$($PriceDecrease_t$) takes the value of 1 when a firm reports a price increase (decrease) and 0 otherwise.

17 The results may be sensitive to the difference in the respective numbers of observations for price increases and decreases. By randomly deselecting observations from the price decrease subsample, we confirm that the main results remain unaltered when the number of observations in each subsample is the same.

their prices are likely to be higher than those of other firms, they refrain from adjusting their prices upwards to avoid the large concomitant profit loss. As for intermediate good cost factors, these do not exhibit such asymmetry and the differences in the coefficients on price increases and decreases are relatively small.[18] This may be the result of cost shocks common to rival firms. In a kinked-demand framework, shifts in demand following changes in relative prices are the source of asymmetric price setting. When firms face a common cost increase, it is unlikely to induce relative price changes, and consequently they do not exhibit asymmetric pricing.

When we add the Taylor dummy variables to control for time-dependent pricing, the above results remain unchanged (see Koga et al (2019) Appendix Tables 1 and 2).

Inflation expectation[19]

Next, we examine the connection between firms' inflation expectations and their price-setting stance. Unlike previous studies exploring the relation between expected price changes in firms' own goods markets and their price setting (eg Boneva et al (2016)), our focus is on how firms' expectations regarding general inflation affect their price setting, and whether these effects vary depending on the direction of the price change and the time horizon. The results are shown in Koga et al (2019), who report the positive association between firms' inflation expectations and their pricing attitude.

5. Extensions

In this section, we further investigate heterogeneity in strategic complementarity in pricing by extending the analysis to include market structure, and examine the effect of demand uncertainty on firms' price setting. The results are not shown here due to space constraints, and are reported in Koga et al (2019).

Strategic complementarity and market structure

The effect of strategic complementarity on pricing may differ across markets. Our conjecture is that when firms' pricing power is stronger, they can set their prices without needing to worry about competitors' prices.[20] We add an interaction term for market share and competitors' prices to examine this hypothesis. The market share of each firm is calculated based on annual sales volume as reported in the survey.

18 The existing studies demonstrate that prices tend to respond faster to input increases than to decreases in various markets (eg Peltzman (2000), Loupias and Sevestre (2013)).

19 Since the first quarter of 2014, The *Tankan* survey has been collecting firms' assessments of the inflation outlook for both general prices and the output prices of their own products or services; it is the former item which we use in this paper. Firms are asked how general prices will have changed relative to current levels one year, three years and five years ahead on an annual basis. They choose from an incremental series of 1% inflation ranges, starting at –3% and going up to +6%. We replace each inflation range with its midpoint and label this variable $EInflation_{i,t}^{l_{YY(s)}}$ (l=1, 3 and 5).

20 Amiti et al (2019) present a model where the price elasticity of demand depends on market share, and argue that strategic complementarity is stronger when a firm's market share is larger.

Our estimation results show that the coefficient on the interaction term for competitors' prices and market share is significantly negative, suggesting that firms with a high market share do not care greatly about their competitors' pricing stance. The estimates suggest that a 10% higher market share means around 3–4 percentage points less impact from competitors' prices on average. As an alternative approach, we replace market share with the market concentration ratio computed using the Herfindahl-Hirschman Index for 636 industry categories, thus capturing the degree of competition at the industry level. Firms in monopolistic industries show a reduced sensitivity to competitors' prices relative to those in competitive industries. In addition to the interaction term, the independent market share and market concentration ratio terms also have negative coefficients, reflecting the fact that firms with higher pricing power or in a monopolistic industry are likely to adjust their prices less often.

Demand uncertainty

We now turn to the question of whether heightened demand uncertainty affects the probability of changing output prices.

We examine the impact of uncertainty on a firm's price-setting behaviour in the manner employed in Bachmann et al (2013) and Bachmann et al (2019). As our data set contains firms' responses to demand changes in both the preceding quarter and the following quarter, we can compute the measure of subjective uncertainty as in their studies. We calculate the difference between the expected demand change in the current period and the realised demand change in the next period, ie the forecast error, and take absolute values. Our interest is in whether firms' price-setting behaviour is affected by uncertainty regarding demand in the next quarter.

Our results show that uncertainty makes price adjustment by firms more likely. This finding is consistent with the existing studies. In addition, we also examine whether demand uncertainty has a significant impact on the responsiveness of firms' price setting to shifts in demand conditions. The results show that when firms face uncertainty the impact of shifts in demand on price setting is reduced. In other words, under uncertainty, they are reluctant to adjust their prices, even when demand conditions change.

Our findings, therefore, broadly confirm the results of the previous studies, but they also extend these to provide new evidence of "wait and see" pricing in the case of demand uncertainty.

6. Conclusions

Using a large panel of firm survey data from Japan, we examine firms' price-setting behaviour. Our paper contributes to the existing literature by providing micro evidence for firms' price setting under a quasi-kinked demand curve. Under such a mechanism, pricing decisions by firms are mutually reinforcing, and firms tend to be cautious about raising their prices. Specifically, we find the following results. First, we find evidence for strategic complementarity in firms' price setting across broad sectors. Second, firms' reactions to their competitors' prices are asymmetric depending on the direction of price adjustment, in accordance with the theoretical predictions of the quasi-kinked demand curve setting. Specifically, they tend to be more responsive to price reductions by competitors than to price rises. Third, we find a positive relationship between the inflation expectations of firms and the probability of them increasing prices. Fourth, the degree of strategic complementarity differs across firms: firms with greater pricing power are likely to care less about competitors' prices. Fifth, heightened demand uncertainty promotes price adjustment by firms, and also mitigates the effect of shifts in demand conditions on the likelihood of price adjustment.

Our findings provide implications for a prolonged deflation that Japan has experienced. Strategic complementarity in pricing is one explanation for firms' cautious stance: firms refrain from increasing their own prices because their competitors are doing the same. This caution may be left behind once some firms start to adjust their prices upwards, as interactive pricing behaviour reinforces the upward adjustment. In addition, our results suggest that encouraging firms to expect higher inflation acts to break them out of their conservatism towards price adjustment.

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Comments on "Strategic complementarity and asymmetric price setting among firms"

Martin Berka[1]

Summary of the paper

Koga et al (2019) is an interesting, well written and timely paper based on data from quarterly *Tankan* surveys covering around 10,000 firms in Japan between 2004 and 2017. Although the paper contains a number of secondary findings, there are five key findings. First, firms' pricing decisions exhibit strategic complementarity, in that they are affected by the pricing decisions of competitor firms. Second, this complementarity is asymmetrically stronger when prices decline, a finding the authors attribute to the existence of kinked demand. Third, higher inflation expectations raise the likelihood of any given firm raising its own price. Fourth, firms with a larger market share exhibit much less sensitivity to competitors' price changes. And finally, firms that report a higher degree of uncertainty delay price changes, an evidence of "wait and see" behaviour. The overall contribution here is one of more details on the state-dependent nature of price changes.

In their well written paper with a thorough literature review, the authors work hard to link the empirical estimations to the theoretical model in Dotsey and King (2005), in which a firm's price relative to the overall price level depends on a number of characteristics of the market environment. Koga et al (2019) further log-linearise the pricing equation of Dotsey and King (2005). Although this is not a structural equation in the sense of mapping prices to some fundamental drivers (instead, the prices depend on costs, labour costs, local demand, and the prices of competitors), Koga et al (2019) use it to derive their empirical model to bring to the data. Because the prices are not observed in the *Tankan* survey, the authors use a limited dependent model (ordered probit).

Comments

In my discussion of the paper, I elaborated on a number of comments. I think the key comment is that the evidence of the asymmetry uncovered in the empirical results of Koga et al (2019) does not need be driven by a kinked demand curve. The authors may want to think more broadly, and in any case beyond the model of Dotsey and King (2005), when trying to interpret their results. But given that the equation is not a structural demand equation, it may just as well be that the "kink" to which the authors attribute the asymmetry is on the supply rather than the demand side. However, the point I want to make here is even broader. The economic literature on the economics of price adjustment is indeed very large, and there will be a number

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of models that could possibly support their empirical findings. For example, Diamond (1971) presents a theoretical framework of price adjustment to a long-run, non-competitive equilibrium. In the model, firms exhibit different market shares, and their pricing decisions depend on both short-term profits and longer-term market share considerations. Each firm *j* maximises

$$\alpha_t^j p_t^j X_t\left(p_t^j\right) + V_t^j\left(\alpha_{t+1}^j\right)$$

where α_t^j is the market share of firm *j* at time *t*, $X_t\left(p_t^j\right)$ is the demand for firm *j*'s product, and $V_t^j\left(\alpha_{t+1}^j\right)$ is the value of future profits, dependent on the future expected market shares of firm *j*. Even with such a simple introduction, it seems possible that the Diamond (1971) model may bear out a number of the empirical regularities discussed in Koga et al (2019). First, if competitors' prices change, Diamond's firm will worry about losing its market share and follow suit. Because this threat is present when prices decline rather than increase, the co-movement should be asymmetrically stronger for the cases of price declines. Second, it is easy to think of a reason why firms worry about market shares less when they are larger.[2] Thus, while I am not formally evaluating the fit of Diamond's 1971 model to Koga et al's data, it seems that the model could capture their first, second and fourth results, without invoking a kinked demand scenario. I think it is likely there are many other models the authors could allude to with similar or better "success" in explaining their findings. It would be a welcome contribution if they discussed a wider family of possible theoretical interpretations of their empirical findings.

Additionally, the paper would benefit from clarifying a number of issues. It would help to understand why a survey of 220,000 firms is distilled to a mere 10,000 observations. Although the *Tankan* survey is reasonably well known, it would still be worth spelling out that it only contains surveys of pricing intentions, rather than surveys of actual prices. Furthermore, it would help to clarify whether the observations in the survey refer to changes from "*t*" to "*t*+1", or "*t*–1" to "*t*". This is not clear from the quotation included from the survey definitions. Furthermore, it is not clear whether the survey records the data for "current" and "forecasted" variables as a single (difference) variable, or as two separate variables.

The study concerns the behaviour of a firm and its competitors. It would therefore seem key to have a well identified measure of competitors. The paper identifies competitors for firm *j* as all other firms in a given industry. While this may be a good description of reality for some industries that are highly traded, possibly in durables, there are many other economic sectors where competition is highly localised. It would be useful to include the location information (eg distance) in the analysis, if available. Furthermore, given that the price adjustment relative to "competitors" is in fact a measure of price adjustment relative to the industrial average, a worry I have is that industry-specific shocks are causing within-industry price dispersion, and this is being misinterpreted as a strategic complementarity.

By the virtue of only containing data on qualitative margins, the study is predisposed to find support for time- rather than state-dependent pricing models.

2 For example, the capacity of competitors to "absorb" switching customers can be limited by the competitors' size, which in many industrial structures is likely to be inversely related to the size of the firm we study. Or, the fraction of customers who are actively engaged in searching for cheaper alternatives is limited, and such a fixed number is of a lesser consequence for a larger firm than for a smaller one. A number of other reasons likely exist.

Thus, it is reassuring that the results are very significant. However, the authors may wish to discuss in more detail the temporal frequency of price adjustment. Given that 74% of all quarterly observations are "no change", a more careless reader may interpret this as "once a year" price adjustment, quite common in time-dependent pricing literature. It would be worth dispelling that concern.

I think it would also be worthwhile to spend more time discussing the size of the asymmetry: the authors find that the complementarity is four times larger when prices are being cut. Furthermore, the asymmetry is only significant when also controlling for market power, although this masks the finding that firms with a higher market power respond less to competitors than those with a smaller market power. I think these findings should be moved out of the extensions, and into the main results. Furthermore, a firm's *own* market share lowers the price change probability – this is a new finding in the literature, and should also make it to the main results, I think.

Finally, I would like to ask the authors to think about linking their findings on "demand uncertainty" to a literature on business cycle turning points. Business cycle turning points are clearly periods of heightened uncertainty of demand, and given that the sample spans the period around the Great Financial Crisis, the data contain a good period to check the robustness of the "wait and see" hypothesis. Presumably, the "wait and see" results should be stronger then than immediately prior to the onset of the crisis.

Let me just conclude by saying that I have learned a lot of interesting and stimulating new facts that seem to give further support to the state-dependent pricing hypothesis in the world's third largest economy. I hope that my comments will provide the authors with some ideas to further improve and strengthen their paper.

I am also grateful to the organisers of the conference for inviting me to their stimulating event.

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Impact of relative price changes and asymmetric adjustments on aggregate inflation: evidence from the Philippines

Joselito R Basilio and Faith Christian Q Cacnio[1]

Abstract

The paper uses disaggregated price data to determine whether the higher moments of the distribution of relative price changes provide information on the adjustments and persistence of aggregate price conditions in the Philippines. It takes into account the changes that occurred in relative price movements between the pre-inflation targeting (ie 1994–2001) and inflation targeting (ie 2002–September 2019) periods. Results indicate that the dispersion of relative price changes and the skewness of their distribution are positively related to movements in short-run inflation. Moreover, price adjustments are observed to be asymmetric, which can have significant effects on short-run inflation.

JEL classification: E3, E31, E52.

Keywords: relative price changes, distribution of price changes, asymmetric price adjustments, inflation, inflation targeting.

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1. Introduction

Classical theory makes a distinction between inflation and relative price changes. Inflation, as Milton Friedman pointed out, is always and everywhere a monetary phenomenon. It is fluctuations in money supply that determine the price level. Relative price changes are determined by real factors such as gyrations in the supply and demand for different goods. Thus, in theory, real price changes should not affect inflation. Friedman emphasised this point when he wrote about the high inflation rates in the early 1970s:

> *"It is essential to distinguish changes in relative prices from changes in absolute prices. The special conditions that drove up the prices of oil and food required purchasers to spend more on them, leaving less to spend on other items. Did that not force other prices to go down or to rise less rapidly than otherwise? Why should the average level of all prices be affected significantly by changes in the prices of some things relative to others?"* (Friedman (1974, p 74))

Accordingly, with an unchanged money stock, relative price adjustments are made through increases in the nominal prices of some goods and decreases in others. Ball and Mankiw (1995) noted that Friedman's analysis implicitly assumes that nominal prices are perfectly flexible. However, this is not often the case in the short run. Frictions, like menu costs, can affect price changes. Firms experiencing price shocks will only change their prices if the desired adjustment is large enough to warrant paying the associated menu cost. Moreover, asymmetries in price adjustments could arise (Ball and Mankiw (1994)). Positive shocks to firms' desired prices are more likely to result in greater adjustments than negative shocks of the same size. This implies that asymmetric adjustments in relative prices could be inflationary in the short run.

Empirical studies in this area (eg Vining and Elwertowski (1976), Fischer (1981), Amano and Macklem (1997)) have looked at the statistical relationship between the higher moments (ie variance and skewness) of the distribution of price changes and inflation.[2] These studies have generally observed that relative price variability is closely associated with fluctuations in aggregate inflation and that short-run movements in inflation are positively related to the skewness of the distribution of relative price changes.[3]

Following this line of inquiry, this paper looks into the link between the distribution of relative price changes and short-run inflation in the Philippines. The paper uses disaggregated price data to determine whether the higher moments of the distribution of price changes provide information on the adjustments and persistence of aggregate domestic price conditions. The empirical exercises yield additional observations after controlling for the other factors that affect headline prices (eg oil prices, rice supply conditions, seasonality and business cycles). One

[2] An earlier work that investigated the movements in individual prices relative to the aggregate price level is that of Mills (1927).

[3] While a number of studies conducted for different countries have found a strong correlation between inflation and its higher moments, these empirical findings have been questioned by the work of Bryan and Cecchetti (1999) and, to some extent, Verbrugge (1999). These authors argued that the observed correlation between inflation and its higher moments is due to small-sample bias. However, Ball and Mankiw (1999) countered that Bryan and Cecchetti's claim is based on their analytical model's departure from the classical model's explanation on the factors that affect the general price level.

particular factor that the paper takes into account in its analysis is the adoption of inflation targeting (IT) in the Philippines in 2002. The paper assesses the changes that may have occurred in relative price movements between the pre-IT (ie 1994–2001) and IT (ie 2002–September 2019) periods.

Our results indicate a link between relative price variability and short-run inflation in the Philippines. High inflation periods, and to a lesser extent, deflationary episodes are associated with higher levels of price change variability. Additionally, the skewness of the distribution of price changes was observed to be positively related to movements in inflation. During periods of rising inflation, a positively skewed distribution suggests that some commodities are experiencing larger price changes relative to the others and these are putting an upward pressure on the general level of prices. The tails or shocks to prices of the different goods and services can be isolated and these can have significant effects on overall inflation under certain conditions. The higher moments of the price distribution can likewise provide some explanation on the observed decline in the sensitivity of short-run inflation to demand pressures. Between the pre-IT and IT periods, the frequency of price changes declined and the duration between price adjustments increased to 1.6 months from 1.4 months. These findings partly explain the low price volatility and stable inflation that the country experienced during the IT period and the observed flattening of the Phillips curve.

The paper is outlined as follows: Section 2 describes the data used and provides some initial observations on relative price changes in the Philippines; Section 3 explores the link between relative price change distribution and short-run inflation in the country; Section 4 looks into the asymmetry of price changes and provides an assessment of its relationship to inflation; and Section 5 concludes.

2. Description of the data and some initial observations

The paper uses the Philippine Statistics Authority's (PSA) disaggregated monthly CPI data (ie three-digit commodity groups)[4] for the period January (M1) 1994–September (M9) 2019. The data set contains 94 items categorised under 11 major commodity groups (Table 1). Some of the CPI items have missing values. These items were either non-existent in the earlier part of the time series (eg mobile phones and phone cards in the early 1990s) or were not included in the CPI basket.

[4] Disaggregated CPIs used in the study are coded with three digits (hence, the reference as three-digit CPI). This is except for the "cereals" CPI. Following the PSA standard of separating rice, corn and other cereals (which are all sub-items of the "cereals" CPI even if "cereals" is already a three-digit CPI) in its statistical table releases, the analysis in this paper likewise treats rice, corn and other cereals as three-digit items.

CPI commodity groups (2012 = 100)		Table 1
		No of 3-digit CPI items
1.	Food and non-alcoholic beverages	13
2.	Alcoholic beverages, tobacco, etc	5
3.	Clothing and footwear	6
4.	Housing, water, electricity, gas and other fuels	8
5.	Furnishings, household equipment and routine maintenance of the house	11
6.	Health	7
7.	Transport	10
8.	Communication	3
9.	Recreation and culture	17
10.	Education	7
11.	Restaurant, miscellaneous goods, and services	7
	ALL ITEMS	94

Source: Philippine Statistics Authority.

The varying patterns of price changes are explored using disaggregated (ie three-digit level) CPI data. Table 2 shows the average frequency of price changes for each commodity group as well as the average duration of the gap between these price movements. These are based on the changes in the prices of the different commodity items that are categorised under each group. Table 2 yields some initial observations about price movements in the Philippines:

1. Prices in the Philippines, on average, changed in about 75.6% of the months in the sample period (ie price increases and decreases). Between 1994 and M9 2019 (ie 308 months), there were 200 months when prices increased and 29 months when they decreased. Price increases, on average, occurred every 1.5 months while price decreases happened every 10.4 months.
2. Increases in prices were seven times more likely to occur than price decreases and almost three times more prevalent than no price change.
3. Prices stayed the same at an average rate of once every four months, or equivalently 24.4% of the entire period.
4. Price decreases, on average, accounted for a relatively small percentage at 12.7% of the total price changes (ie price increases and decreases).
5. Among the commodity groups, food and non-alcoholic beverages, alcoholic beverages and tobacco, housing, utilities, gas and other fuels, and restaurant and miscellaneous goods experienced higher rates of price changes relative to other commodity groups. Quite the opposite is the case for education and communication, where the frequencies of price changes are substantially lower than the others.

A significant policy shift that occurred over the 1994–M9 2019 sample period is the adoption of inflation targeting (IT) as the framework for monetary policy in the Philippines in 2002.[5] Empirical studies (eg Guinigundo (2017)) have observed that changes occurred in the country's inflation dynamics following the adoption of IT. Inflation persistence gradually declined as the inflation process shifted from being

[5] In January 2002, the BSP formally adopted inflation targeting as the framework for monetary policy. Based on initial assessments, the adoption of inflation targeting helped the country sustain a favourable inflation performance over the medium term (Guinigundo (2005)).

backward-looking to more forward-looking. Bangko Sentral ng Pilipinas (BSP) managed to keep inflation within target, leading market agents to adopt a more forward-looking view in their assessment of current inflation. With increased monetary policy credibility, expected inflation started to weigh more in the pricing decisions of firms and consumers.

To assess the potential impact of IT on price changes, we divide our sample period into 1994–2001 and 2002–M9 2019 and compare the movements in relative prices between these periods. In Table 2, prices (on all items), on average, are shown to have changed less (ie price increases and decreases) during the IT period (73.2%) relative to the pre-IT period (79.8%). The frequency of price increases declined significantly in the IT period (63.8%) compared to the 1994–2001 period (70.8%). Moreover, the proportion of price decreases and no price changes increased in the 2002–M9 2019 period.

Among the commodity groups, food and non-alcoholic beverages, which has the largest weight in the CPI basket, as well as alcoholic beverages, tobacco, etc, clothing and footwear and education experienced higher shares of price increases in the IT period relative to the pre-IT period. Most of the items in the food and non-alcoholic beverages commodity group (eg rice, other cereals, fish and seafood, milk, cheese and eggs) had a higher frequency of price increases in the 2002–M9 2019 period compared to the 1994–2001 period. Price increases in food items, particularly of agricultural commodities, are for the most part due to weather-related disturbances that cause lower supply and disruptions in the supply chain. The incremental increase in the tax rates of alcoholic beverages and tobacco products resulted in significant adjustments in the prices of alcoholic beverages and tobacco products starting in 2014. Meanwhile, lower proportions of price increases were observed for the commodity groups of housing, water, electricity, gas and other fuels, restaurant, miscellaneous goods and services and transport in the IT period relative to the pre-IT period.

The duration of the gaps between price increases was, on average, 1.4 months over the 1994–2001 period. This lengthened to 1.6 months in the 2002–M9 2019 period. Within the commodity groups, there was a notable lengthening in the duration between price changes for transport, communication, recreation and culture and education in the 2002–M9 2019 period.

The observed decline in the frequency of price changes and the lengthening of the duration between price adjustments correspond to a period of lower average inflation in the economy. Inflation declined from an average of 7.6% between 1994 and 2001 to 3.8% in the 2002–M9 2019 period. The rates of price change for the different commodity groups likewise declined in the later period.

Frequency and duration of price changes: 1994–M9 2019, 1994–2001 and 2002–M9 2019, CPI commodity groups

2012 = 100

Table 2

	1994–M9 2019			1994–2001 (Pre-IT)			2002–M9 2019 (IT)		
Total no of months	308			95			213		
	Increase	Decrease	No change	Increase	Decrease	No change	Increase	Decrease	No change
Frequency of price change (in no of months)									
All items	200	29	7473	63	8	18	136	20	57
Food and non-alcoholic beverages	215	71	23	63	22	10	152	48	14
Alcoholic beverages, tobacco, etc	270	15	23	80	8	8	191	7	15
Clothing and footwear	257	13	38	78	6	11	179	8	27
Housing, water, electricity, gas and other fuels	235	46	27	75	10	9	160	35	18
Furnishings, household equipment	244	16	48	76	7	12	168	9	36
Health	253	13	42	80	3	12	173	9	31
Transport	150	45	113	45	9	41	97	29	88
Communication	79	61	169	40	13	43	52	42	119
Recreation and culture	177	26	106	64	8	23	109	20	84
Education	55	7	188	11	2	24	44	6	163
Restaurants, miscellaneous goods and services	262	10	37	82	4	9	170	8	35
Share to total number of periods (in per cent)									
All items	66.0	9.6	24.4	70.8	9.0	20.2	63.8	9.4	26.8
Food and non-alcoholic beverages	69.8	23.1	7.5	66.3	23.2	10.6	71.4	22.5	6.6
Alcoholic beverages, tobacco, etc	87.7	4.9	7.4	84.2	8.4	8.4	89.7	3.4	7.2
Clothing and footwear	83.4	4.4	12.3	82.3	6.1	11.6	84.0	3.6	12.6
Housing, water, electricity, gas and other fuels	76.3	14.9	8.8	79.4	11.0	9.6	75.1	16.2	8.5
Furnishings, household equipment	79.2	5.2	15.5	80.0	7.4	12.9	78.9	4.2	16.9
Health	82.1	4.1	13.7	84.4	4.2	10.5	81.2	4.4	14.4
Transport	48.7	14.6	36.7	47.4	9.5	43.2	45.5	13.6	41.3
Communication	25.6	19.8	54.9	42.1	13.7	45.3	24.4	19.7	55.9
Recreation and culture	57.5	8.4	34.4	67.4	8.4	24.2	51.2	9.3	39.4
Education	22.0	2.8	75.2	29.7	5.4	64.9	20.7	2.6	76.5
Restaurants, miscellaneous goods and services	85.1	3.1	12.0	86.3	4.2	9.5	79.8	3.7	16.3

Duration of gap between price changes (median average, in no of months)									
All items	1.5	10.4	4.1	1.4	11.1	4.9	1.6	10.7	3.7
Food and non-alcoholic beverages	1.5	3.9	14.0	1.5	4.5	8.6	1.5	3.6	17.8
Alcoholic beverages, tobacco, etc	1.1	25.7	11.8	1.2	15.4	13.6	1.1	53.3	11.8
Clothing and footwear	1.2	24.5	7.3	1.2	23.8	9.5	1.2	23.5	7.7
Housing, water, electricity, gas and other fuels	1.3	9.4	15.8	1.3	7.9	13.6	1.3	9.1	20.7
Furnishings, household equipment	1.3	18.1	5.9	1.3	16.3	7.6	1.4	18.2	4.7
Health	1.3	34.2	5.9	1.2	19.0	7.7	1.3	30.4	5.3
Transport	2.6	16.4	1.8	1.7	13.6	2.6	2.6	16.3	1.9
Communication	3.3	7.2	2.2	1.5	19.0	8.6	4.4	7.6	2.0
Recreation and culture	1.7	13.7	3.0	1.5	10.0	5.1	1.7	13.7	2.8
Education	5.4	44.3	1.3	4.6	20.6	1.4	5.1	50.7	1.3
Restaurants, miscellaneous goods and services	1.3	33.0	6.3	1.2	39.6	8.6	1.3	42.6	5.8

Details may not add up to total due to rounding.

Sources: Philippine Statistics Authority; authors' calculations.

3. Relative price change distribution and aggregate inflation

Following the work of Vining and Elwertowski (1976), we determine the link between relative price changes and short-run inflation in the Philippines. We do this by looking at the distribution of price adjustments and the corresponding shape of the distribution of the price changes. The starting point is the calculation of the frequency of price adjustments as well as non-adjustments. It is then followed by the estimation and analysis of the corresponding (changing) shape of the distribution over time in terms of the second and third moments (ie standard deviation and skewness).

To estimate the frequency and magnitude of price changes for each *j*th three-digit level disaggregated (CPI) item, we use the following equations:[6,7]

Frequency of price changes: $$F_j = \frac{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUM_{ijt}}{\sum_{i=1}^{n_j} \sum_{t=2}^{t} DEN_{ijt}} \quad (1)$$

Frequency of price increases: $$F_j^+ = \frac{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUMUP_{ijt}}{\sum_{i=1}^{n_j} \sum_{t=2}^{t} DEN_{ijt}} \quad (2)$$[8]

Average price increase in per cent $$\overline{\triangle}_j^+ = \frac{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUMUP_{ijt} (\ln P_{ijt} - \ln P_{ij,t-1})}{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUMUP_{ijt}}$$

(3)

where the price available at *t* is given as $DEN_{ijt} = 1$ if P_{ijt} and $P_{ij,t-1}$ are observed in *t*; 0 otherwise. Price change at *t* is defined to be: $NUM_{ijt} = 1$ if $P_{ijt} \neq P_{ij,t-1}$; 0 otherwise.

To distinguish between price increases and decreases, the former is set as:

$NUMUP_{ijt} = 1$ if $P_{ijt} > P_{ij,t-1}$; 0 otherwise.
Meanwhile, a price decrease at *t* is set as: $NUMDW_{ijt} = 1$ if $P_{ijt} < P_{ij,t-1}$; 0 otherwise.

The distribution of price changes was derived for each month. The shape of these distributions, in turn, allowed us to generate the monthly values of the higher moments of the distribution.

Previous studies that examined the relationship between relative price changes and short-run inflation have highlighted two important observations: (i) relative price variability is closely associated with fluctuations in aggregate inflation; and (ii) short-run movements in inflation are positively related to the skewness of the distribution of relative price changes. We determine whether these observations hold for the case of the Philippines.

6 These equations and corresponding descriptions are from Abenoja and Basilio (2018).

7 The *i*th refers to the geographical location for which the three-digit CPI data is available for each region in the country. It is not included in the actual computations.

8 One can also estimate the frequency of price decreases and the average price decrease by changing the $NUMUP_{ijt}$ with $NUMDW_{ijt}$ in equations (2) and (3).

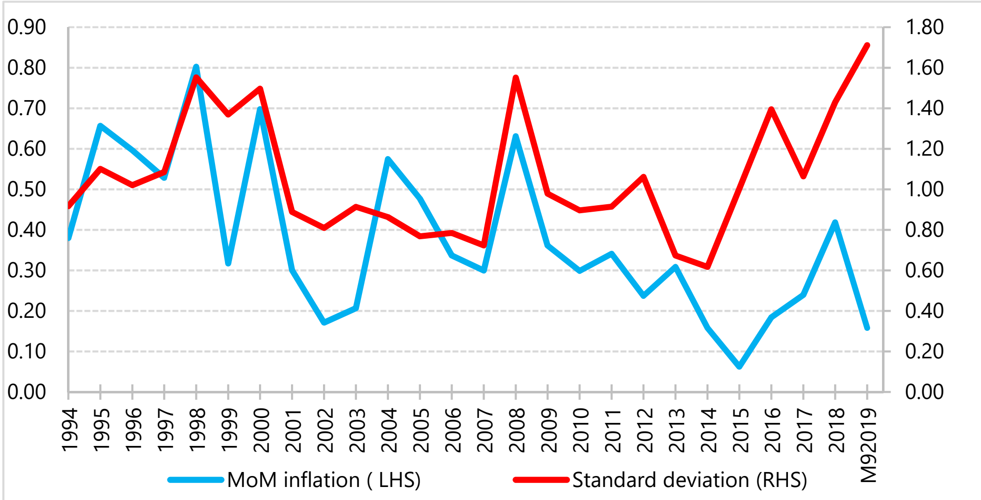
First observation: The variability of relative price changes is closely associated with movements in short-run aggregate inflation.

In Figure 1, we present a comparison of the inflation rate and the variance of relative prices. Inflation rate is the month-on-month (MoM) rate of increase in the CPI, while the variance of relative prices is the standard deviation of the rates of change (MoM) of the 94 individual commodities of the CPI under consideration. Figure 1 reveals that high variability of relative price changes in the Philippines points to rising levels of inflation. Higher relative price change variability implies a higher frequency of price changes in the economy. To a lesser degree, higher dispersion of relative prices can also indicate a period of deflation (2016). The observed peaks in inflation (ie 1998, 2000, 2018) were attributed to supply side shocks. Still reeling from the impact of the 1997 Asian crisis, the Philippines experienced poor weather conditions and drought in 1998 which adversely affected its agricultural harvest. This led to double-digit food inflation during the year. In 2000, rising oil prices and higher electricity rates drove up non-food inflation. Higher food and energy prices likewise caused the increase in the inflation rate in 2018. Meanwhile, in 2015 and 2016, low international oil prices and ample food supply largely contributed to the decline in the rate of inflation. An important observation that appears in Figure 1 is the low level of relative price variability between 2001 and 2007 despite high inflation rates in 2004–06 due to supply side shocks. A possible explanation for this is that the supply shocks that occurred during this period triggered second-round effects (ie increases in transportation fares, higher utility charges, adjustments in minimum wages across the country) that led to higher inflation. Moreover, the national government implemented tax reform measures in 2005 and 2006. In 2005, the value added tax (VAT) exemptions for several industries, including power, electricity, air and sea transport, were lifted. Energy and oil companies were allowed to pass on the 10% VAT to their consumers. The following year, in 2006, the national government increased the VAT rate for goods and services from 10% to 12%. These developments contributed to a permanent increase in the prices of most of the goods and services in the economy.

Inflation rate and standard deviation of relative prices

Month-on-month, in per cent

Figure 1



Sources: Philippine Statistics Authority; authors' calculations.

Figure 1 also reflects the observed decline in the volatility of inflation following the adoption of IT. Inflation volatility declined from an average of 2.5% in the 1994–2001 period to 2.0% in the period 2002–M9 2019. With the exception of 2008 and 2016, relative price variability remained relatively low over the 2002–18 period.

Second observation: Short-run movements in inflation are positively related to the skewness of the distribution of relative price changes (Ball and Mankiw (1994, 1995)).

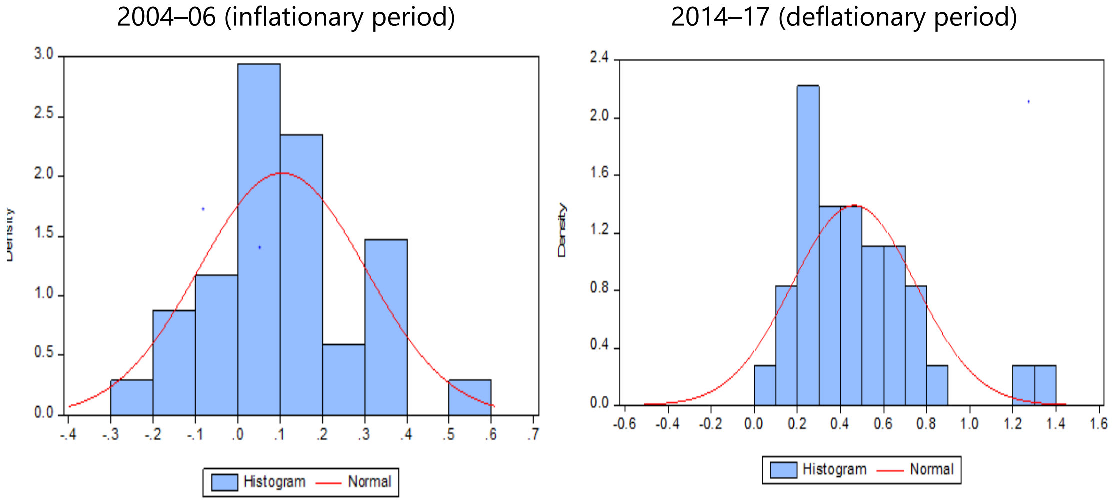
An assessment of the distribution of the price changes in the Philippines shows that it is positively skewed. This relates to the finding in Table 2, which shows a higher frequency of price increases in the country relative to price decreases or no price changes. Moreover, a positively skewed distribution of relative price changes signifies that price increases in certain commodities could be large enough to result in inflationary pressures. The distribution became less positively skewed and more symmetrical in the 2002–M9 2019 period relative to the 1994–2001 period.

We look at two particular periods in our data to see whether the skewness of the distribution of relative price changes provides information on the direction of aggregate inflation. The first period is 2004–06 (inflationary period) and the second is from the latter part of 2014 to early 2017 (deflationary episode). In Figure 2, the distribution of price changes during the period of relatively high inflation is shown to be positively skewed while the period with declining prices has a slightly negatively skewed distribution.

During periods of rising inflation, a positively skewed distribution implies that some commodities are experiencing larger price changes which outweigh possible price decreases (or no price movements) in the other commodities. The converse holds during periods of declining prices. In the Philippines, price pressures have always been attributed to supply side shocks, particularly to food, oil and energy items.

Distribution of relative prices

Figure 2

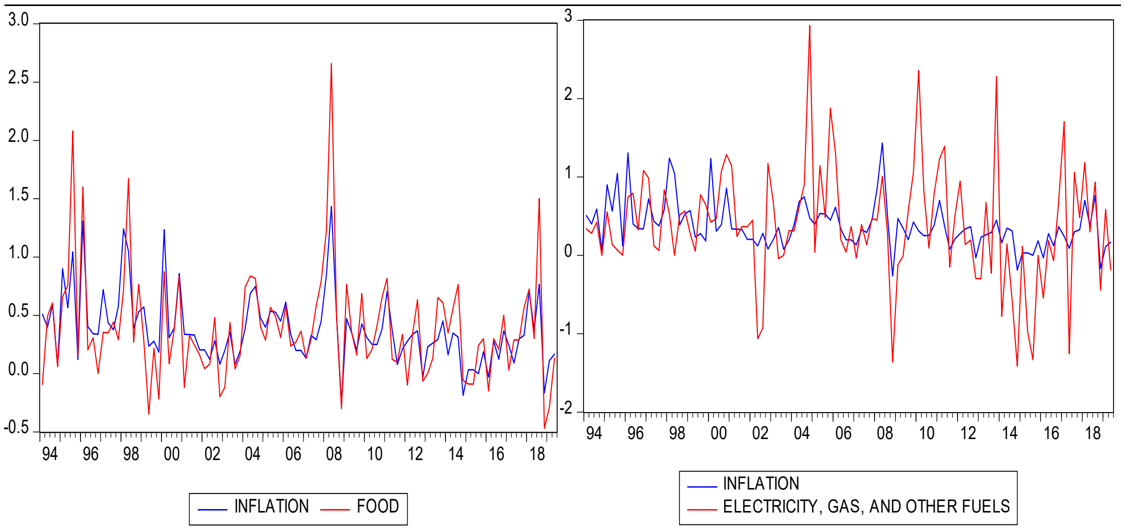


Sources: Philippine Statistics Authority; authors' calculations.

Figure 3 plots inflation against the price changes in food, oil and energy.[9] The graphs show that, during periods of inflation (deflation), price changes in the food, oil and energy items are larger (smaller) than average inflation. This signifies that price changes in these commodities pull up (down) the general level of prices in the economy.

Inflation and relative price changes in selected commodities

Figure 3



Sources: Philippine Statistics Authority; authors' calculations.

[9] The sub-commodity group food is comprised of rice, corn, flour, cereals, bread, pasta, other bakery products, meat, fish and seafood, milk, cheese, eggs, oils and fats, fruits, vegetables, sugar, jam, honey, chocolate and confectionery, coffee, tea, cocoa, mineral water, soft drinks, fruit and vegetable juices.

4. The asymmetry in the distribution of price changes and inflation

We looked at the asymmetry of the distribution of the price changes in the Philippines to see how it relates to the positive relationship observed for inflation and the relative price variability and skewness of price changes. Data on the distribution of the price changes and inflation and the availability of the monthly series of variances and skewness allowed us to analyse the corresponding asymmetry of the distribution. Consistent with the methodology used in Ball and Mankiw (1995), we estimated the asymmetry indices for each month over the sample period of January 1994 to September 2019. The output of the estimation produces the index for measuring the asymmetry of the distribution of disaggregated price changes/inflation for each month. Figure 4 plots two measures of asymmetry. The upper panel is based on the distributions over time of the month-on-month inflation figures for the 94 disaggregated CPI items. The lower panel measures asymmetry based on the distributions of month-on-month difference in the CPI level for the same cross section of CPI items.

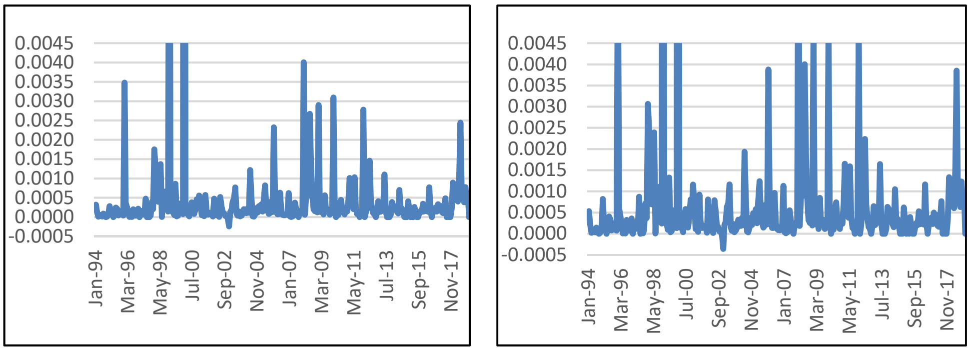
Estimates of the asymmetry index based on distributions of inflation data and CPI price level changes show consistency of results. For example, both charts in Figure 4 generally have the same peaks. Periods with large tails usually coincided with one-off periods where specific CPI items exhibited irregular spikes in inflation. This was the case for the peaks in January 1996 (due to the implementation of the expanded Value Added Tax (E-VAT) law (Jurado (2017)), January 1999 (period of higher food prices related to the occurrence of El Niño/La Niña and a period of depreciating currency during the Asian crisis), January 2000 (higher prices of crude oil amid a newly deregulated downstream oil industry and major change in the base year and components of the CPI), January 2006 (implementation of the Reformed VAT or RVAT Law and hikes in world crude oil prices), January 2008 (sustained increase in global oil prices that peaked at an all-time high in mid-2008), January 2010 (El Niño period), January 2012 (implementation of the "Sin Tax" Law), and January 2018 (implementation of a comprehensive tax reform package and the eventual rise in global oil prices and domestic rice supply issues).

The measures of asymmetry capture large movements in both relative prices and aggregate inflation. Moreover, applying various cutoff points for estimating asymmetry did not reduce the high correlation of the two measures. The use, therefore, of either of the two measures and the changing of the cutoff points should not affect the regressions and the corresponding interpretation of the regression results. Further analysis also yields the observation that the values of the asymmetry index are generally positive. This signifies an upward trend in price indices (ie positive inflation) over time, despite downward adjustments for some items for given periods.

Estimated asymmetry indices, January 1994–September 2019

2012 = 100 Figure 4

Asymmetry index based on MoM inflation | Asymmetry index based on MoM CPI price level adjustments



Sources: Philippine Statistics Authority; authors' calculations.

4.1 Inflation and price asymmetry: some regressions

We use the asymmetry indices generated in the previous section to assess short- run inflation dynamics in the Philippines based on the following specification:

$$\pi_t = \alpha \quad (4) \quad \pi_{t-1} + \theta_\sigma \sigma_t + \theta_\kappa \kappa_t + \theta_{Asym} Asym_t + \textstyle\sum_{\forall i} \theta_i exo_{t,i} + \varepsilon_t$$

where

π_t	=	short-run inflation (month-on-month CPI change, in per cent)
α	=	constant for the regression
π_{t-1}	=	lagged π_t
σ_t	=	standard deviation of the distribution of disaggregated price changes over time *t*
κ_t	=	skewness of the distribution of disaggregated price changes for each period *t*
$Asym_t$	=	indicator variable for asymmetry (in index points), for each period *t*
$exo_{t,i}$	=	the other *i*th exogenous variable that significantly affects aggregate inflation (eg seasonality, output gap, oil prices, rice prices)
ε_t	=	period *t* error term for the regression

θ_π, θ_σ, θ_κ, θ_{Asym}, θ_i are the respective coefficients of the explanatory variables.

Results show that both measures of asymmetry have significant effects on short-run inflation (Table 3). This observation holds even if the moments of the distribution (ie standard deviation and skewness variables) and other exogenous variables (eg oil and rice prices) are added into the regression. Such a result confirms the standard theoretical basis for equation (4), which considers the addition of the moment variables and other exogenous variables as regressors for short-run inflation.

Skewness, which measures the symmetry of the whole distribution and not just the tails, was found to be significant when included in the regressions (columns 3 and 4). The standard deviation of the distribution of disaggregated price changes was significant in the regression, even with the inclusion of the rice price variable (column 4). However, the standard deviation was not significant when the regression adds oil prices (column 4). This was not surprising considering that the large volatilities or dispersions across a cross section of CPI inflation occurred at the same time as that of (or even as a result of) higher global oil prices. There is an expected degree of correlation between variables for standard deviation and global oil prices. This also explains why world oil prices tend to lack statistical significance in their effect on inflation.

Regression results using alternative measures of asymmetry

Table 3

	Dependent variable: inflation			
	(1)	(2)	(3)	(4)
Constant	0.2317	0.2307	0.1171	0.1251
	(0.0272)	(0.0250)	(0.3498)	(0.0311)
Lagged inflation	0.2344	0.2315	0.2282	0.2358
	(0.0458)	(0.0424)	(0.0420)	(0.0412)
Asym_Adj[1]	67.2726		53.2560	60.7434
	(5.5462)		(6.7343)	(5.6416)
Asym_Inf[2]		23.8389		
		(1.5961)		
Standard deviation			0.0666	0.0166
			(0.0218)	(0.0194)
Skewness			0.0310	0.0292
			(0.0050)	(0.0045)
Crude oil price				0.7261
				(0.4339)
Rice prices				7.5041
				(1.2454)
Adj R-squared	0.3515	0.4450	0.4576	0.5932
DW[3]	1.7829	1.6870	1.8091	1.8822

Standard errors are in parentheses. Number of observations: 307.
[1] Asym_Adj pertains to the asymmetry index that is measured based on the distributions of CPI level changes. [2] Asym_Inf is based on CPI inflation or percentage changes. [3] DW = Durbin-Watson test.

The adjusted R-squared ranges between 0.3515 and 0.5932. This is fairly comparable with the results in Ball and Mankiw (1995). Durbin-Watson tests indicate the rejection of serial correlation.

Regressions of aggregate inflation lead to an analysis of the Phillips curve when adding some output or employment variables into the equation. Table 4 presents the regression results of equation (4) with output gap and unemployment as additional explanatory variables. The output gap is not positive or statistically significant, even as various proxy variables (real GDP growth) or transformations (ie in levels, per cent or logs) are used. The coefficient of unemployment is significant but of the wrong sign (positive). We note that while these results may not be consistent with those observed for advanced economies (eg the United States in Ball and Mankiw (1995) or Canada in Amano and Macklem (1997)), the empirical results are consistent with the Philippine experience. In the 1980s and until the late 1990s, the Philippines registered low growth rates and high average inflation. However, starting in the late 2000s, the country started to achieve higher growth rates that were accompanied by generally lower and more stable inflation.

The lessening sensitivity of prices to real economic activity (ie flattening of the Phillips curve) in the Philippines has been attributed to the adoption of IT (Guinigundo (2017)). Nonetheless, the IMF (2006) pointed out that while improved monetary policy credibility can account for a large part of the decline in the sensitivity of prices, more than half of it is accounted for by other factors, including global factors. In the case of the Philippines, increased trade openness is cited as an important factor that led to lower frequency of price changes and the lengthening of the duration between price adjustments in the economy.[10] Prices responded sluggishly to domestic demand pressures given increased trade and investment flows. Strong international competition constrained firms and businesses from increasing prices even when demand rose.

In Table 4, lagged inflation is positively significant across all the regressions, indicating the degree of persistence of past inflation performance. Meanwhile, the asymmetry variable is consistently observed to have a positive and significant effect across all regressions. The asymmetry variable, as a transformation of (or the differential between) the mass of the tails of the distribution of price changes, is by construction an indicator of the outliers in a cross section of price changes of the disaggregated CPI items. This outlier effect is one of the possible information content items of the asymmetry index that is generally independent of skewness and standard deviation.

Consistent with the results of the regressions, the standard deviation of the relative-price changes loses statistical significance when world oil price is included in the model as an exogenous variable. Similar to the interpretation of Amano and Macklem (1997), the information contained in oil prices of the distribution appears to be redundant. It duplicates the independent effects of the dispersion of disaggregated price changes on aggregate price conditions.

Without world oil prices in the Phillips curve regressions (columns 1, 2 and 3), the coefficient of standard deviation is positive and significant. This finding is consistent with the results of the menu cost framework in Ball and Mankiw (1994, 1995). In contrast to the results of standard Phillips curve-menu cost models, the effects of the interaction of skewness and standard deviation on aggregate inflation is found to be statistically insignificant. This lack of significance is the reason for dropping the interaction variable in the other regressions. The interaction variable relatively provides no new information under the described setup.

[10] Starting in the 1980s and until the 2000s, the Philippines undertook key trade reforms and policies geared towards liberalising and improving the domestic economy's competitiveness. Between the 1980s and early 1990s, the Philippines pursued trade reform programmes that substantially reduced import tariffs. The country also acceded to the AFTA-CEPT (1993) and GATT-WTO (1995). In the 2000s, the Philippines pursued trade facilitation through regional and bilateral free trade agreements.

Philips curve equations

Table 4

Dependent variable: inflation					
	(1)	(2)	(3)	(4)	(5)
Constant	–0.1460	0.1460	–0.2161	–0.1355	0.1448
	(0.0830)	(0.0388)	(0.0804)	(0.0739)	(0.0346)
Lagged inflation	0.3186	0.2160	0.3020	0.2424	0.2286
	(0.0531)	(0.0426)	(0.0518)	(0.0500)	(0.0415)
Output gap	–7.18E-07	–2.25E-06			–1.50E-06
	(1.27e-06)	(1.34e-06)			(1.17e-06)
Unemployment	0.0310		0.0338	0.0311	
	(0.0101)		(0.0098)	(0.0092)	
Asym_Adj[1]	160.3721	52.8072	146.8666	147.1478	60.3677
	(20.1422)	(6.7550)	(20.0018)	(18.2080)	(5.6419)
Standard deviation	0.0121	0.0646	0.0583	–0.0002	0.0156
	(0.0230)	(0.0219)	(0.0273)	(0.2093)	(0.0193)
Skewness	0.0276	0.0316	0.0457	0.0276	0.0295
	(0.0045)	(0.0051)	(0.0076)	(0.0041)	(0.0045)
Standard deviation x Skewness			–0.0112		
			(0.0034)		
Log difference of world oil price				0.3176	0.7676
				(0.4039)	(0.4346)
Log difference of domestic rice price				7.2227	7.4640
				(1.1603)	(1.2442)
Adj R^2	0.5069	0.4606	0.5279	0.5955	0.5943
DW[2]	1.9637	1.7952	2.0008	2.0672	1.8751
No of observations	192	304	195	192	257

Standard errors are in parentheses.

[1] Asym_Adj pertains to Asymmetry index as measured based on the distributions of CPI level changes. [2] DW = Durbin-Watson test.

4.2 Subsample splitting and possible structural changes between the pre-IT (1994–2001) and IT (2002–19) periods

Comparing between two periods (ie pre-OPEC, 1949–69 and OPEC, 1970–89), Ball and Mankiw (1995) observed the general stability of coefficients derived from the regression of aggregate inflation against asymmetry, skewness and standard deviation. The subsample regressions excluded food and energy variables as these were not statistically significant. Rather than these traditional measures of supply shocks, the asymmetry index turned out to be better at measuring supply shocks over time (Ball and Mankiw (1995)).

As in Section 2 of this paper, we split the sample period into the pre-IT (1994–2001) and IT (2002–19) periods. The former is a much shorter subsample than the latter.

Corresponding changes in the magnitude and significance of the coefficients can be observed from the regression results for each subsample (Table 5). Lagged inflation, for instance, is not significant for the pre-IT period (columns 1 and 3) but becomes significant during the IT period (columns 2 and 4). Using lags as proxy for expected inflation (Amano and Macklem (1997)), this finding reflects the important role of inflation expectations during the IT period. Inference shows an absence of evidence on the role of expectations or persistence during the pre-IT period, given the lack of significance of the coefficients of lagged inflation. The output gap variable shows similar degrees of insignificance for both subsamples. Furthermore, its coefficients changed sign from positive (pre-IT period) to negative (IT period).

The asymmetry index is significant for both periods, while seeing an increase in the magnitude of its effects (coefficient) during the IT period. The coefficients of the asymmetry index in columns 2 and 4 are about three to four times the size of its coefficients in columns 1 and 3, respectively. Meanwhile, it is difficult to draw conclusions about the role of standard deviation (ie dispersion of inflation in a cross section of the disaggregated CPI items) as its coefficients are generally not significant (for columns 1 and 2). In a Phillips curve setting (columns 3 and 4), the standard deviation was significant for the pre-IT period (column 3) but not significant for the IT period (column 4). The larger dispersions during the pre-IT period (when compared to the IT period) could partly explain its greater role in explaining aggregate inflation during that period.

Similar to the regression results presented in Tables 4 and 5, the skewness is significant and shows some stability in both the significance and the magnitude of its coefficients. World oil prices (in logs) were not significant in the various regressions (columns 1 and 2). Meanwhile, domestic rice prices were not significant during the pre-IT period but became significant during the IT period. Rice prices were more volatile during the pre-IT years, which partly explains the observed lower significance during this period. Overall, there is subsample stability of coefficients for asymmetry, skewness and lagged inflation. As stated in Ball and Mankiw (1995), the role of traditional indicators of supply shock (food and energy) in the regressions of aggregate inflation may relatively "matter only because they induce asymmetry in the distribution of price changes".

Subsample stability and possible structural changes
(pre-IT and IT periods)

Table 5

	Regressions with oil and rice prices			
	Pre-IT (1994–2001) (1)	IT (2002–M9 2019) (2)	Pre-IT (1994–2001) (3)[1]	IT (2002–M9 2019) (4)[1]
Constant	0.2554	0.0893	0.2352	0.0736
	(0.0900)	(0.0309)	(0.0780)	(0.0384)
Lagged inflation	0.0736	0.2691	0.0505	0.3455
	(0.0825)	(0.0485)	(0.0719)	(0.0511)
Output gap			4.45E-07	–4.81E-07
			(3.88e-06)	(1.24e-06)
Asym_Adj[2]	53.5526	148.4848	38.1913	161.5308
	(10.8063)	(18.2205)	(10.5398)	(19.9755)
Standard deviation	0.0069	(0.0059)	0.0951	0.0075
	(0.0482)	(0.0205)	(0.0426)	(0.0223)
Skewness	0.0574	0.0274	0.0546	0.0272
	(0.0187)	(0.0040)	(0.0139)	(0.0044)
Log difference of world oil price	1.9991 (1.4446)	0.3759 (0.3992)		
Log difference of domestic rice price	4.1156 (5.4819)	7.1419 (1.1516)		
Adj R^2	0.6839	0.5680	0.5177	0.4809
DW[3]	1.6804	2.0515	1.9299	1.9729
No of observations	47	209	94	209

Standard errors are in parentheses.
[1] An alternative Phillips curve regression (ie columns 3 and 4) was tested with unemployment as a substitute for the output gap. The unemployment variable was significant but only during the IT period. [2] Asym_Adj pertains to Asymmetry index as measured based on the distributions of CPI level changes. [3] DW = Durbin-Watson test.

5. Concluding thoughts

In summary, this paper looked into the link between the distribution of relative price changes and short-run inflation in the Philippines. Disaggregated price data was used to determine whether the higher moments of the distribution of price changes provide information on the adjustments and persistence of aggregate domestic price conditions.

The analysis in this paper yielded important observations for prices in the Philippines. Some of these are in keeping with empirical observations in other countries. First, there is a close association between the variability of relative price changes and short-run inflation in the Philippines. Episodes of high inflation were characterised by higher levels of price change variability. To a lesser extent, higher price dispersion was also associated with a deflationary period. Second, the skewness of the distribution of price changes was observed to be positively related to the movements in inflation. During periods of rising inflation, a positively skewed distribution signifies that some commodities are experiencing larger price changes relative to the others and these are putting upward pressure on the general level of prices. Third, the tails or shocks to prices of different goods and services can be isolated. Evidence from the Phillips curve regressions shows that these can have significant effects on overall inflation under certain conditions. Fourth, the higher moments of the price distribution can provide some explanation on the observed decline in the sensitivity of short-run inflation to demand pressures.

The assessment of the impact of IT on relative price movements showed that, between the pre-IT and IT periods, the frequency of price changes declined. Additionally, the duration of the gap between price adjustments increased from 1.4 months in the pre-IT period to 1.6 months in the IT period. These findings partly explain the decline in average inflation and low price volatility that the country experienced over the past 17 years. In a regression analysis, splitting the sample between the pre-IT and IT periods indicated that the asymmetry and skewness of the distribution of price changes affect aggregate inflation in the Philippines.

Going forward, further work in this area still needs to be done. For example, the reasons behind the asymmetry in price adjustments (ie motivation for menu costs) could be further explored. It would also be interesting to use a more disaggregated data set, ie the five-digit CPI items, in order to further account for the heterogeneity and dispersion of price changes within markets (of the same product/sector) and across different markets.

The observations from this paper contribute to a better understanding of inflation dynamics in the Philippines, which is important for monetary policy. The finding that relative price changes can be inflationary in the short run could complicate the conduct of monetary policy. While monetary policy can influence inflation, it cannot really affect relative price changes. Monetary policy likewise needs to take into account the asymmetries in price adjustments which could also cause inflationary pressures in the short run. These are important considerations that policymakers need to keep in mind.

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Comments on "Impact of relative price changes and asymmetric adjustments on aggregate inflation: evidence from the Philippines"

By Renée Fry-McKibbin[1]

1. Overview of the paper

The objective of Basilio and Cacnio's paper is to examine the link between the distribution of relative price (rp) changes and short-run inflation (π) for the Philippines between 1994 and 2019. The authors follow the approach of Ball and Mankiw (1995), who focus on higher-order moments such as skewness in the distribution of relative price shocks and the implications for inflation. The presence of higher-order moments leads to asymmetric price adjustments in response to shocks compared to the case with normally distributed relative price changes. This issue is pertinent when considering the effects of supply shocks, which are likely to affect the distribution of relative prices most. Supply shocks are often a cause of inflation in the Philippines, particularly in the food and oil sectors, and are likely to become more prevalent in the face of climate change. Hence, monetary policy depends on knowing the nature of the shock and how changes in relative prices might affect inflation.

Using disaggregated monthly data for 94 items that make up the CPI, Basilio and Cacnio use the moments of the distribution of the relative price changes to show that the distribution of elements of the CPI data is non-normal. They then calculate an asymmetry index of the distribution of shocks to relative prices. The measure looks at the difference between the mass in the upper and lower tails of the distribution of shocks. It is then used in a regression model of inflation. The authors also perform a simple regression of inflation on the standard deviation and skewness of the distribution of relative prices, oil prices and rice prices. They find that there is a relationship between the distribution of relative price changes and inflation in the short run. Importantly, they find that the tails matter.

2. Comments

The Ball and Mankiw result implies that there are significant implications of the interactions of first, second and third moments of relative price changes and inflation. Presumably, we are also interested in the timing of the changes between skewness in relative prices and the level of inflation. The authors could draw upon the literature on financial market crises and contagion to identify if the relationship changes, and timing of the distribution of relative prices on the inflation rate. This suggests a cokurtosis-based test that could be used to examine changes in the interaction

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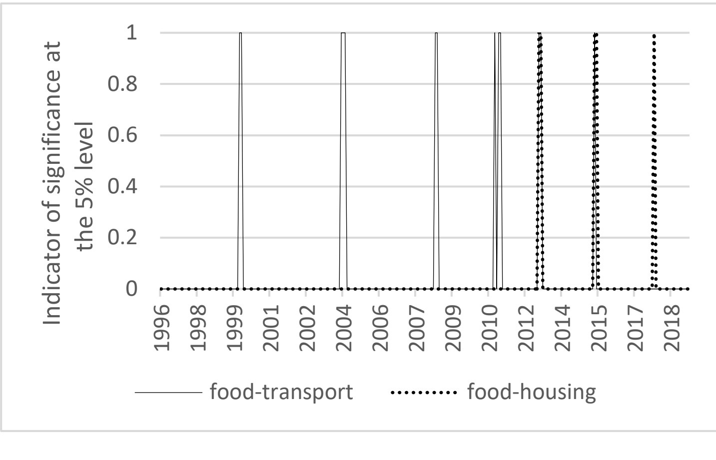
between skewness in the distribution of relative prices and the level of the inflation rate.

This idea was first developed by Forbes and Rigobon (2002), who identify contagion through changes in the correlation of two asset markets. They compare the correlation in a normal period x denoted $\rho_x(rp^1\pi^1)$ with the correlation in a crisis period y denoted $\rho_y(rp^1\pi^1)$ with an adjustment for changes in heteroscedasticity and with a source of the change identified. The asset markets in the Forbes and Rigobon test may be replaced by the relative price and inflation data in the comparison periods and the source of the change in the relationship is relative prices. Fry et al (2010) extended this idea to examine changes in coskewness comparing the joint distribution of the volatility of variable one and the level of variable two in period $x, \varphi_x(rp^2\pi^1)$ and period $y, \varphi_y(rp^2\pi^1)$. The pertinent test in this family of distributions is to examine the change in cokurtosis developed in Fry-McKibbin and Hsiao (2016). In the relative price-inflation context this test could be applied to determine changes in the joint distribution of the skewness of relative prices with the inflation rate across time, where cokurtosis is denoted $\vartheta_x(rp^3\pi^1)$ in the first period and $\vartheta_y(rp^3\pi^1)$ in the second period.

To illustrate the application of these simple tests to the relative price and inflation joint distribution, two indices of relative price changes are calculated using: i) the ratio of food and non-alcoholic beverages to transport (food-transport); and ii) the ratio food and non-alcoholic beverages to housing, water, electricity, gas and other fuels (food-housing). Figure 1 shows an indicator of the significance of the change in the joint distribution of the cokurtosis between the relative price variable and inflation $\vartheta(rp^3\pi^1)$. The tests are conducted using a rolling sample of a window of 30 months in period x and 30 months in period y. The figure shows that significant changes in the relationship between relative prices and inflation are not that common. There are 13 instances where there is a change in the joint distribution between food-transport and inflation, and five instances between food-housing and inflation. The results show that there are only two instances of changes in the joint distribution in the pre-inflation targeting period. These occurred in January and February of 2000. The relationship between skewness in relative prices and inflation is more prevalent in the inflation targeting period.

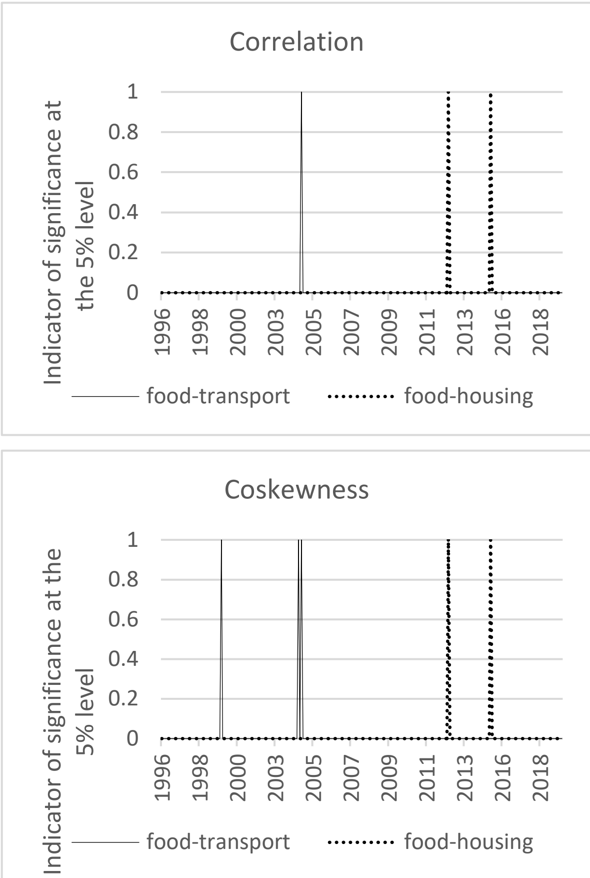
Figure 1



Indicators of the significance of cokurtosis $\vartheta(rp^3\pi^1)$ change tests between relative prices and inflation in the Philippines, January 1994–July 2019. The tests compare $\vartheta_x(rp^3\pi^1)$ and $\varphi_y(rp^3\pi^1)$ over 30-day rolling windows. The relative prices are food prices to transport and food prices to housing.

Figure 2 presents the indicators of significance for the changes in the joint distribution of relative prices to inflation through the correlation and coskewness. The joint distribution based on correlation changes only once for the food-transport case, and twice for the food-housing case. Changes in coskewness only occur three times for the food-transport case, and twice for the food-housing case. The only significant change occurring through coskewness before the inflation targeting period began is in January 2000, corresponding to the change in cokurtosis at this time. There are no changes in correlation before the inflation targeting period. Further work on determining the nature of the shocks that occurred in conjunction with the significant changes would be useful in informing the conduct of monetary policy in response to future shocks. These tests can be modified to be one-sided tests (the current version is a two-sided test), which may further illuminate the relationships between relative prices and inflation joint distributions.

Figure 2



Indicators of the significance of the correlation $\rho(rp^1\pi^1)$ and coskewness $\vartheta(rp^3\pi^1)$ change tests between relative prices and inflation in the Philippines, January 1994–July 2019. The tests compare the statistics over 30-day rolling windows. The relative prices are food prices to transport and food prices to housing.

3. Conclusion

Basilio and Cacnio confirm the results of Ball and Mankiw (1995) that asymmetries in relative price movements are significant for inflation by using disaggregated price data for the Philippines. Interestingly, they find that in the inflation targeting period, the frequency of price changes was lower, and the duration between price adjustments was longer than in the pre-inflation targeting period. In comparison, the results of the higher-order comoment change tests show that significant comoment changes occur infrequently in the latter period, and hardly at all during the pre-inflation targeting period. This suggests that less frequent changes in price adjustments are likely to correspond with a change in the relative price-inflation joint distribution.

Basilio and Cacnio could further explore the robustness of the relationship between asymmetry in relative prices and the inflation rate before and after the introduction of inflation targeting in January 2002. In its current form the paper does not distinguish the change in regime, and the models used are estimated over a relatively long time horizon. It would be useful to know whether relative price shocks affect inflation differently in the inflation targeting period compared to the former monetary policy regime. Perhaps this information could be used to improve the responses of monetary policy to shocks in the presence of relative price changes.

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