The Great Disconnect: The Decoupling of Wage and Price Inflation in Japan*

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Abstract

We take some well-known observations about the structure of the Japanese labor market and add new evidence about how it has evolved to study inflation in Japan. Our key finding is that labor market dynamics shifted after 1998 so that measures of labor market tightness and wages weakened noticeably. This change was accompanied in a break between wages and prices, so wage inflation is no longer an important determinant of price inflation.

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1. Introduction

Japan's persistently low inflation is puzzling in two respects. On the one hand, the Bank of Japan (BoJ) since 2013 has engaged in unprecedented monetary stimulus. This began with a massive asset purchase program (concentrated on Japanese government bonds), which was followed by driving short term interest rates below zero and ultimately targeting the yield on the 10 year bond rate. Despite these efforts, and the conventional view that a determined central bank can raise the price level, inflation has remained below the BoJ's target of 2% per year.

On the other hand, the labor market in Japan has been relatively tight. There is no broadly accepted estimate of the natural rate of unemployment in Japan. However, speeches by BoJ officials often indicate that they believe that at an unemployment rate of about 3% the only people unemployed are in that situation because of search frictions. With unemployment at 2.8 percent, one would expect to see rising wage and price inflation (even if the monetary policy stance were neutral). Likewise, standard estimates of potential growth in Japan are about 0.5 percent. The economy has been consistently above this level since early 2013. Thus, adherents of the Phillip's curve might also find the subdued inflation puzzling.

In this paper we explore the role of Japan's dual labor market in driving the low inflation outcomes in Japan. Economists have long studied the differences in employment conditions between full-time regular workers and part-time employees in Japan. As the Bank of Japan (2017) notes, there has been a steady rise in part-time workers and their wages are more sensitive to market conditions than full-time workers. However, the level of part-time workers' wages is far below those for full time workers. This change in composition has depressed firm's wage bills. We investigate the effects of this factor on inflation outcomes in Japan.

Our analysis begins by pointing out that there have been three inflation regimes since the end of Japan's high growth era in the early 1970s. From the early 1970s until the second global oil shock in 1980, inflation was high and somewhat volatile. As we show below, if this episode is included in any econometric exercise, the conclusions about inflation dynamics are driven by the wide differences between this episode and the subsequent ones. Thus, we exclude the 1970s from our analysis.

After 1980, inflation is both lower and more stable than in 1970s. Inflation also gradually trended down before rising a bit and then settling around zero by the middle of the 1990s. This pattern holds for both headline measures of inflation and those that exclude food and energy prices. We show that the behavior of wages and prices in this transition period differs from the subsequent period.

The third period begins around 1998. While ability to statistically distinguish the exact break point is limited, there are several good reasons to impose a break in 1998. One is that the Bank of Japan gains its legal independence then. A second is that the acute phase of the banking Japanese banking crisis begins in late 1997. A third, most importantly, is that past research has

found that a number of other structural relations in the economy changed around this time (Hamada, Kashyap, and Weinstein (2011)).[1] Thus, our analysis also separates the pre and post-1998 eras.

Our remaining analysis proceeds in two steps. We begin by marshalling a set of facts about the way wages are set and have behaved in Japan over the last two decades. Most of these observations are familiar to Japanese labor economists, but their importance for inflation has received less attention.

Three important facts stand out. The first is that wages for part-time and full-time workers have evolved very differently since 1997. Full-time wages have stagnated, while part-time wages have trended upwards. Second, these patterns partly reflect different sensitivity of the two types of wages to labor market conditions. Part-time wages tend to rise when the labor market is tight, but full-time wages are much less responsive to overall conditions. Third, the combination of the rising percentage of part-time workers in the overall labor force along with the differences in the level of wages for the two types of workers has meant that the average level of wages has been flat or slightly falling for nearly 20 years. The average wage also appears to be disconnected from the tightness of the labor market (even though the part-time workers' wages are responsive to tightness).

The second step is to investigate the connection between wages and prices. The break in the behavior of wages also carries over to prices in several respects. Prior to 1998, wage and price inflation both were non-stationary; as in many countries, the rates of inflation were very persistent and did not seem to revert to any particular level. Importantly, prices and wages were co-integrated, meaning that their trend behavior was linked. The dynamic relation was such that movements in wages predict changes in prices.

After 1998, wage and price inflation are both stationary. The levels of prices and wages do not share a common trend, and the dynamic link between wages and prices also breaks. Wage movements no longer predict price movements. The connections between oil prices and headline prices also shifted.

The remainder of the paper is separated into five sections. The next one briefly reviews the literature on inflation in Japan. Broadly, past work can be separated into explanations that begin with a Phillips curve approach and then modify it to account for particular factors in Japan. A second set of analyses are closer to ours and put the structure of the labor market at the center of the story. Section 3 presents the evidence that supports our contention that it is wise to analyze inflation separately before and after 1998. We include both our own statistical analysis and review relevant related prior work that argues for the important structural break around 1998.

[1] For instance, one widely studied question is the connection between monetary policy and the economy. Many studies have found a large shift at some point during the 1990s, although they differ in the variables they examine and the exact timing of the break. See, for example, Arai and Hoshi (2006), Kimura, Kobayahi, Muranaga, and Ugai (2003), Kuttner and Posen (2001), and Miyao (2000).

The fourth section of the paper covers our summary of the labor market facts. We establish the three observations described above and explain the important roles of the 1997/1998 banking crisis and the global financial crisis in creating persistent imbalances in the market for full-time workers. Section 5 connects prices and wages. We show how the trends in wages and prices changed in 1998 and how the trend breaks also correspond to shifts in the dynamic relations between the two series. Section 6 presents our conclusions.

2. Literature Review

There is an enormous literature on the Japanese inflation experience over the last two decades. Since the BoJ was given formal independence in 1998, many of these papers have focused on the operating tactics and strategies of the BoJ. Many observers criticized the BoJ for being ambivalent about whether it should fight the ongoing deflation and if so whether it in fact had the necessary tools.

One influential analysis by Bernanke (2000) argued that the BoJ had been insufficiently aggressive in using the tools at its disposal to combat deflation. In a 2003 speech, Bernanke (2003) followed up this analysis and gave a number of suggestions about steps that could be taken to raise the price level in Japan. Svensson (2003) offered what has come to be known as the "foolproof" way to overcome deflation by relying on monetary induced changes in the exchange rate to eventually drive prices up.

These diagnoses were in part a reaction to the BoJ's contention that deflation was occurring for reasons outside of its control. For example, in January 2002 BoJ Governor Hayami gave a speech describing recent BoJ policy decisions (Hayami (2002)). One key passage was (with emphasis included by us):

> "We changed the main operating target from call rates to a quantitative indicator of liquidity, that is, the outstanding balance of current accounts at the Bank of Japan, and we have substantially increased the balance.
>
> Because this was a very drastic easing measure and was unprecedented in the history of central banking worldwide, we had to consider various issues very carefully before deciding it. *Above all, it was uncertain whether and how expansion of liquidity provision in a situation where the rate is virtually zero percent would influence the economy. Needless to say, the mechanism through which such an expansion affects the economy has not been demonstrated either theoretically or empirically.*"

As late as 2011, the Bank of Japan was still making these kinds of arguments. For instance, Governor Shirakawa gave a speech (with a nearly identical title to Hayami's) in which he claimed (Shirakawa (2011)):

> "The prolonged, albeit moderate, deflation since the end of the 1990s cannot be explained only by short-term and cyclical factors. A more fundamental cause is the long-term downtrend in the growth potential of Japan's economy. When the growth rate continues to decline for a protracted period, people's expectations for future income growth are reduced and firms and households restrain their spending. As a result, downward pressure on prices continues (Chart 8).
>
> It is sometimes said that deflation will be solved only if the central bank provides liquidity more aggressively. Provision of ample liquidity is important, of course, but this alone does not solve the problem of deflation."

These arguments were met with continued criticisms from external observers. See, for example, Ito and Mishkin (2006) for a detailed assessment of the first few years of the BOJ independence and IMF (2012) and Kuttner, Iwaisako and Posen (2015) for subsequent critiques.

Yet, starting in 2013, when Governor Kuroda took the helm at the BoJ, these critiques no longer apply. The BoJ has become very aggressive and implemented many of the ideas advocated by its critics. It has stopped questioning its responsibility for inflation outcomes and has disavowed the idea that it is impossible to end deflation. The recent literature, therefore, has shifted to try to explain why, even with extremely accommodative monetary policy, inflation in Japan has remained below 2 percent.

One strand of research takes a Phillips curve approach and seeks to relate inflation to measures of slack in the economy and expectations. Bernanke (2017) surveys much of this work and adds his own conjectures about the reasons for persistently low levels of Japanese inflation. The conventional view is that the sensitivity of inflation to measures of slack have declined in many advanced economies including Japan. The Bank of Japan (2017) gives a specific example of this argument, contrasting the estimated slope of the Phillips curve before and after 2013 and concluding that the curve is much flatter after 2013.

This reduced sensitivity is supplemented with the view that expectations for inflation in Japan are unanchored. Kuroda (2017) argues that expectations in Japan are much more adaptive or backward looking in other countries, so that factors such as oil price swings that change realized inflation can be particularly important in Japan. Bernanke (2017) adds that one reason why expectations in Japan may not be closely pegged to the BoJ's inflation target is the long-period when inflation undershot the target. In this sense, the BOJ may still be suffering from the past performance despite the change in leadership and philosophy.

Not everyone accepts the idea that the Phillips curve is broken in Japan. Gagnon (2017) argues instead that potential output and full-employment have been chronically mis-measured in Japan. As he notes, standard measures of potential output are constructed under the assumption that over the sample being considered, output is both above and below potential. If instead, the economy was consistently depressed so that potential output was never reached (and the

economy was never at full employment), then standard statistical approaches to deduce potential are doomed. Gagnon argues that this was likely the case in Japan between 1995 and 2016, so the more puzzling issue is why there was not more deflation. He conjectures that downward wage rigidity prevented an acceleration of deflation. He predicts that as the economy continues to recover, it will soon pass the level of full-employment and inflation will appear when it does.

The leading alternative to the Phillips curve approach focuses on the structure of the labor market in Japan. An excellent collection of the essays on this topic is Genda (2017). The 16 chapters in this volume explore a variety of reasons why wages have stagnated in the presence of a tight labor market. Three common themes emerge from these essays. One is that the dual labor market is important and that full-time workers and part-time workers are paid and treated very differently. The increasing share of part-time workers is highlighted in many of the chapters. A second theme is the prevalence of downward wage rigidity and the extent to which firms hesitate to cut wages. One essay by Yamamoto and Kuroda notes that there are many firms that never cut wages over a ten year period. These same firms also fail to raise wages when the economy improves.

A third consideration is the supply side of the labor market. One aspect of labor supply is the steady demographic shift in Japan. As the society has aged, many elderly people continue to want to work past the mandatory retirement age, and have stayed in the labor market often by taking part-time positions. This trend has perhaps been supplemented by a collapse in training within firms. On-the-job-training was once an important factor in raising skill-levels and the decline in this practice has reduced the supply of certain types of high-skilled workers. In cases where there is a mismatch between the hiring needs of firms and the supply of suitable workers, wage increases may not be very effective in clearing the market.

Our hypothesized explanation is very much in line with the arguments found in Genda (2017). While none of the essays specifically match our story, we see our story as a convex combination of the different pieces of evidence in the book. The main difference in our work relative to these essays is to explicitly connect inflation developments to the wage developments.

3. Structural Breaks in Japanese Inflation

Figure 1 shows the three main inflation indicators that we examine in this paper, from 1971 to 2017. The solid green line shows our preferred measure of wage growth which is calculated as the average monthly contractual earnings (including overtime pay) divided by the average number of hours worked in the month for workers at establishments with 30 or more employees. The source of the data is *Monthly Labour Survey* (http://www.mhlw.go.jp/english/database/db-l/monthly-labour.html). The dashed red line shows the so-called "core core" inflation rate which relates to the consumer price index excluding food and energy. The dotted blue line pertains to the headline inflation rate as computed from the total consumer price index. For both of the CPI

measures we adjusted the series to remove the mechanical effects of the three increases in the consumption tax on April 1 in 1989, 1997, and 2014. To make the picture more readable, and eliminate seasonality, we compute the four quarter change in the natural logarithm of each series.

Simple inspection of the figure reveals the three main points that are important for our subsequent analysis. First, the period from the start of the sample through 1980 looks qualitatively different than the rest of the data. This period contains the end of the "high growth era" when Japan's economy averaged more than 9 percent growth, and the two spikes in inflation that accompanied the disruptions in global oil markets in 1974/75 and 1980. Inflation, even excluding food and energy, is much higher and more volatile than the rest of the sample. Both our measures of inflation, and wage growth, rise sharply and then fall during this period.

The episode of the 1970s is so different from what followed after the 1980s that any econometric exercises to examine the connections between the variables will be driven by the events of the 1970s (even though they are not representative of the rest of the sample). There are several ways to see this. For example, Figure 2 shows the scatterplot of the headline inflation rate against the wage inflation for three sample periods: 1971q1 to 2017q3 in Panel A, 1971q1 to 1980q4 in Panel B, and 1981q1 to 2017q3 in Panel C. The fitted regression line is also shown in each scatterplot and the equation for the regression line is given at the bottom of each panel. Comparing Panel B to Panel C confirms how different the experience in the 1970s was in both in the levels of the two variables, but also in their co-movement. When we estimate a simple regression using the entire sample (Panel A), the slope of the fitted line is basically determined what happened in the 1970s; put differently, the slope of the lines in panels A and B are nearly identical, even though panel C shows that the correlation between wage in price inflation after 1980 was lower than before. Although a simple t-test does not reject the hypothesis that the slopes in Panel B and Panel C are equal, a Chow test for structural break strongly rejects the hypothesis.

We do not believe the experience from the 1970s is relevant for our question about why inflation has been so low for the last two decades. So for the remainder of the paper we begin the analysis in 1981.

Starting in 1981 inflation begins falling and declines gradually through 1988 before rising a bit. For the entire time through 1997 inflation sits below five percent. Headline inflation moves more than the core-core measure, but the two variables track each other closely. Wages also move along with the price measures, but are more volatile. However, all three measures fall to about zero by 1998.

Starting in 1998 both price and wage changes hover around zero. All three measures become even more stable than in the prior sample. All three indicators sit consistently in the range of plus or minus 3 percent; the only exception is for the headline inflation rate in the middle of 2014

when, right after the three percent increase in the VAT in April, the reported levels of inflation were slightly above 3%.

To confirm these visual impression that the periods before and after 1998 are different, we conduct two formal statistical tests. First, we used Augmented Dickey Fuller (ADF) tests to assess the degree of integration for the underlying series graphed in Figure 1. The gradual declines in the price changes seen over the period from 1981 to 1998 are such that the test cannot reject the hypotheses that the headline price and wage level series are integrated of order 2, i.e. that the inflation rates are non-stationary. The pre-1998 test for the core-core inflation rate suggests that it is also close to nonstationary.[2] In contrast, after 1998 for both price inflation series and the wage inflation indicator, the test fails to reject stationarity.

In some of our subsequent analysis, we will include an oil price series. Our preferred measure is the standard West Texas Intermediary price of oil converted into yen. For reference, the table also shows the order of integration for this series to be one for both sub-samples. So the degree of stationarity for this series did not change before and after 1998.

As a second set of tests, we estimated Box-Jenkins style time series models for each of the three series and tested for the stability of the models. Because of pronounced degrees of seasonality we include quarterly dummies in these models and these variables are highly significant. We also include three dummy variables that are equal to 1 in the quarters where the consumption tax was raised (1989 Q2, 1997 Q2, and 2014 Q2). Table 2 consists of three panels showing the best fitting models for each of the three series. The models were selected to fit the 1981 to 1997 sample as parsimoniously as possible. The model was then re-estimated over the post-1998 sample to investigate whether the dynamics of the series were stable.

For the headline CPI inflation rate, the best fitting model had no autoregressive terms and lagged moving average terms at quarters two through five. For the core-core inflation rate, the best fitting model also had no autoregressive terms, and lagged moving average terms at quarters two and four. The wage series was modeled as having two lagged autoregressive terms, at quarters one and two, and a lagged moving average term at quarter 4.

We draw three conclusions from these estimates. First, for the two price inflation series, there is only a moderate degree of persistence. While the unit root tests suggested that headline inflation series was I(1), the models that fit best did not use any autoregressive terms. As Cecchetti et al (2017) observe in many countries, a relatively good model for inflation assumes that the first difference of inflation follows a first-order moving average process. The Japanese price series dynamics seem rather different. Second, for both of the price inflation series, the levels of persistence fall after 1998. The moving average coefficients are both smaller in magnitude and

[2] The t-statistic of the coefficient on lagged dependent variable in the ADF test equation for the core-core inflation rate for 1981q1-1997q4 period is -3.63, which is higher than the critical value at 5% level (-3.48) in absolute value, suggesting the inflation rate is stationary (so the level is I(1)), but not as high as the critical value at 1% significance level (-4.10).

have a lower degree of statistical significance. For the headline series, one of the coefficients flips sign. Both price inflation measures rise around the time of the consumption tax increases, although the coefficient estimates are sometimes statistically insignificant.

Third, for the wage series, the dynamics are very different. There is a strong amount of negative first-order serial correlation especially after 1998. This negative serial correlation is sufficiently strong that the models fit quite well, with R^2 around 90 percent. There is no special movement in the wage variable around the dates of the consumption tax increases.

Formal tests for coefficient stability show that for all three series, the coefficients across the two samples are significantly different. For example, a Chow test that supposes a break date of 1998 rejects the equality of coefficients at levels of significance below 1 percent. Since the consumption tax dummies are insignificant in the wage inflation equation, we dropped those when conducting the Chow test.

Alternatively, a Quandt-Andrews test allows for a potential break at any date (other than those that are too close to the beginning or the end of the sample period). Since a Quandt-Andrews test involves estimating the model for various subsamples, including the consumption tax and seasonal dummies in the price inflation equations becomes cumbersome. Thus, we first estimated the model for the entire sample including both quarterly dummies and the tax dummies, and then created a new dependent variable by subtracting the part that can be explained by the quarterly dummies and the tax dummies. We applied Quandt-Andrews test by regressing the new variable on just the constant and the MA terms. We consider a potential break in the middle 70% of the sample (1992Q2 to 2006Q4). This test also strongly rejects the null hypothesis that there were no break points. For headline inflation, the break date that is mostly likely is 1993Q4. For core-core inflation, the most likely break was 1998Q1. For wage inflation, the Quant-Andrews test selects 1999Q2 as the most likely break point.

Combining all of this direct evidence, with the aforementioned analysis in Hamada, Kashyap and Weinstein suggesting that a number of other macro relations changed around 1998, leads us to examine separately the pre and post 1998 data for the remainder of the paper. Our focus now shifts to trying to understand what might be responsible for the shift.

4. Wages and the Labor Market

The structure of the labor market began to shift following the Japanese banking crisis. Figure 3 shows annual data for the average hourly wages in Japan. The figure shows the wage rate for full-time and part-time workers, along with the average for the economy that accounts for both (the green solid line). To calculate the wage series, we first convert the indices of various types of earnings and hours worked at establishments with 30 or more employees into the numbers measured in yen or hours benchmarked against average number for 2012. Then, the wage series

plotted here are calculated by dividing average contractual cash earnings (including overtime) by monthly total hours worked (including overtime) for each group of workers. The data are taken from the *Monthly Labour Survey*. Note that the left axis is the scale for wage series for full-time workers and all workers and different from the scale on the right axis that measures wage for part-time workers. The average level of wages peaked in 1999 and did not exceed that level until 2017.

The components that make up the average have behaved very differently. The wage for regular workers has been remarkably steady sitting around 2100 yen per hour and only slightly increased during the last 20 years. The part-time wage has risen by about 15 percent from below 1000 yen to about 1150 yen per hour between 1997 and 2017. The part-time wage also has been more sensitive to improved business conditions since 2013.

Figure 4 shows that the main reason for the flat average wages despite rising wages for part-time jobs and for full-time jobs is the changing composition of the workforce. We use the seasonally adjusted employment index for each type of job at establishments with 30 or more employees in *Monthly Labour Survey* and convert it into actual number using the ratio of the index value to the actual value at December 2012. Here we show the annual average of each series. The number of full-time workers is measured on the left axis while the level of part-time workers is measured on the right axis. The number of full-time workers peaked in 1994, dips a bit afterwards, before it began a precipitous decline in 1998 that continued through 2004. The number of part-time workers had been relatively stable until 1996 and then began a climb that continued over the next 20 years. The replacement of high cost regular employees with much less expensive part-time workers is a major reason why average wages in 2017 are almost the same as in 1997.

The last section documented changes of dynamics for both prices and wages. For wages, the reason for the changing dynamics is specifically attributable to the interaction of the sharp recession that began in 1997 and the dual labor market. Unfortunately, data on wages for part-time and regular workers in *Monthly Labour Survey* is only available starting in 1993. This prevents us from being able to directly show that there was a break in the link between the labor market conditions and the wages for full-time as well as part-time workers. Consequently, we proceed indirectly, building up evidence in three steps.

First, using data on average wages, which are available since 1970, we show that the connection between average wages and labor market tightness changed around 1997. We do this by augmenting the ARMA model for the wage inflation that we considered in the last section with measures of labor market tightness (or slack) and checking whether the response of wage to labor market conditions changed before and after 1997.

This analysis is presented in Table 3. Here we expand on the ARMAX models considered in the last section, by adding two measures of labor market tightness or slack. One is the ratio of active job openings to job applicants (AJOAR in the table). A large value of AJOAR means that there

are many job openings compared to the number of applicants and thus implies tight labor market condition. As our second proxy we consider the standard unemployment rate.

We use the same AR(1), AR(2), MA(4) specification that we used in the last section for the average wage for all workers. Panel A of Table 3 uses AJOAR as the measure of labor market tightness. The panel shows the estimation results for two different sample periods: 1981-1997 and 1998-2017. For both samples, one period lag of AJOAR is statistically significant, but the size of the coefficient in the later period is less than one-third of the earlier period. Thus, the average wage does respond to the labor market condition in the expected direction in both periods, but the magnitude of the effect drops substantially after 1998.

We obtain a similar conclusion from Panel B, which uses the unemployment rate as the measure of labor market slack. The expected sign of the coefficient on the unemployment rate is negative: high unemployment tends to depress wages. However, the strength of the effect drops noticeably after 1998, again falling to about 1/3 of the pre-1998 estimate. These estimates imply that all else equal, between 1981 and 1997, when unemployment rose by 0.5 percentage points in one quarter, inflation would rise by about 1 percentage point in the next quarter. A similar increase after 1998 would predict about a 0.3 percentage point increase in the next quarter.

The flattening of the wage Phillips curve in Japan has been noted in various studies including the recent paper by Muto and Shintani (2017). So the estimates in Table 3 can be read as confirming the observation that average wages now seem only weakly connected to shortages or excesses in the labor market.

To investigate what might be behind this change, we turn to data on market conditions for full-time (aka regular) workers. We do this by looking at how the proportion of enterprises that report excess full-time workers changed over time using data from *Rodo Keizai Doko Chosa (Survey of Labour Economy Trend)* conducted by the Ministry of Health, Labour and Welfare in February, May, August, and November of each year.

Figure 5 shows the proportion of enterprises that express they have excessive full-time workers. This variable is clearly cyclical, and the level of excess workers jumps in 1998 and stays elevated till 2004, even though there is a partial economic recovery in 2000. By 2005, the excess employment indicator is back to the 1997 level and it remains there for 2 more years, before spiking again during the global financial crisis. Looking back at Figure 4, it is clear that the periods of excess correspond to times when full-time employment falls. Our contention is that this adjustment is incomplete, and if firms had not hoarded workers the number of full-time employees would have fallen much more sharply and presumably would have recovered more quickly.

This hunch leads to our third piece of evidence that explores whether measures of slack have been equally important for full-time and part-time workers. To do this we estimate ARMAX models for wages separately for full-time and part-time workers. Because the wages data for

full-time and part-time jobs separately are available only from 1993, we conduct the analysis for the post-1998 period only.

These results are shown separately in panels A and B of Table 4. The first panel shows several alternative model specifications for full-time workers. Models 1 and 2 start from a second order autoregressive specification, augmented with seasonal dummy variables, and adds respectively the lagged job opening to applicant rate (for the entire economy) and the (economy wide) unemployment rate. Both proxies for slack are weakly correlated with wage growth for full-time workers. The unemployment rate is border-line statistically significant, though the size of the coefficient is very small. A one half of a percentage point increase in the unemployment rate would predict a 0.2 percentage point decrease in full-time worker wages in the next quarter.

Models 3 and 4, shown in the lower part of the panel, add estimates of the job opening to job applicant ratios for full-time and part-time workers as proxies for slack. In model 3, only the job opening rate for full-time workers is included, while in model 4 both measures are used. In either case, the connection between slack and wage inflation is weak. Thus, all four models in panel A are consistent with the view that wages for full-time workers after 1998 were divorced from the tightness of the labor market. Notice also that the R^2 for all 4 models are around 94%, so these models do a good job of tracking wage inflation.

The regressions in panel B investigate whether the same patterns are true for part-time workers. We estimate the same type of specifications as in panel A, with the main difference being that we include a three period lagged moving average error as an additional control. In models 1 and 2 the slack variables are lags of the economy-wide job opening rate and the unemployment rate, respectively. In each case, the slack measures are highly statistically significant. The magnitude of the unemployment coefficient implies that a one half percentage point increase in unemployment leads to about a 0.45 percentage point increase in wages in the next quarter.

Models 3 and 4 use the measures related to job openings to job applicants for part-time and full-time workers as indicators of slack. In model 3, we find that the part-time job opening to job applicant rate is weakly correlated with part-time wage inflation. Interestingly, the number of full-time job openings to full-time job applicants is a strongly statistically significant predictor of wage inflation for part-time workers -- so much so that it completely drives out the part-time job opening rate.

Given that these labor market variables are not predictive of what happens to full-time wages, and the evidence in section 3, we are led to the following account of the data. The market for full-time workers was often out of balance after 1998. This meant that labor market tightness did not influence full-time wages. Instead, the marginal equilibrating factor in the labor market occurs around part-time workers. When the market is relatively tight for full-time workers, wages respond for part-time workers. This is not so surprising since many people would prefer to work full-time, so to keep them in part-time positions their wages much rise.

We also find it interesting that the R^2 for the various models shown in panel B are lower than in Panel A. Predicting what is happening to part-time wages is harder than predicting full-time wages.

5. Wages and Prices

If the trend behavior of prices is driven by trend movements in wages, then the change in the wage dynamics could explain why Japanese inflation continues to be low despite recent economic recovery following the massive monetary accommodation. So we investigate this hypothesis. Because we found breaks in both price dynamics and wage dynamics in 1998, we investigate the co-movements between prices and wages separately before and after 1998. We also consider oil price to be another essential variable that influences the prices and wages in Japan. Indeed BOJ (2016) lists the development of oil price as one of the most important reasons why inflation has been stagnant despite expansive monetary policy.

Figure 6 plots the levels of headline CPI and core-core CPI with the index for the average wage for all workers. The structural break around 1998 that we found above is apparent in this figure as well. Before 1998, both wage and prices trended up, but they became stagnant or started to fall after 1998. The declining trend seems stronger for the core-core CPI than for the headline CPI or the average wage.

To simplify the discussion below, we use the following notation for the remainder of the analysis.

PA: Log of the headline CPI
WA: Log of average wage for all workers
PO: Log of the oil price in yen

For 1981-1997 period, the Augmented Dickey Fuller Tests in Table 1 suggest that PA and WA are I(2) while PO is I(1). To assess trend movements, we first consider a system of three endogenous variables: ΔPA, ΔWA, and PO. Assuming each of the three variables is I(1), we apply a Johansen test of co-integration to the system. To address strong seasonality of the data and the impact of the VAT increases, we include three centered quarterly dummies and the first difference of the tax dummies that we introduced earlier.[3]

Table 5 reports the results of the Johansen cointegration test. For the number of lags for the VAR system, we chose 2, which is the number suggested by the AIC criteria for the unrestricted

[3] A centered quarterly dummy takes 3/4 in a certain quarter and -1/4 in the other three quarters, so that the four quarterly dummies sum up to zero. We use centered dummies instead of the standard 0-1 seasonal dummies to avoid contaminating the inference about the possible trend of the endogenous variable as suggested by Johansen (1995, p.84). We first difference the tax dummies because they predict movements in the level of prices, so to analyze the change in prices we need to take differences in the dummies.

system. The trace test suggests 2 cointegrating relations, while the maximum eigenvalue test suggests 1 cointegrating relation. Accordingly, we estimate two error correction models: one assuming a single cointegrating relation and the other assuming 2 cointegrating relations.

Table 6 shows the estimates from the error correction model of (ΔPA, ΔWA, PO) assuming 1 cointegrating relation. The estimated cointegration equation is given by:

$$\Delta PA - 0.8054*\Delta WA - 0.0017*PO + 0.0165$$
$$(-5.66) \qquad (-1.68)$$

The numbers in parentheses below the coefficient estimates are t-statistics. We normalize the coefficient on ΔPA to be 1. The coefficient on the co-integrating equation in the table suggests that equilibrium is restored via headline inflation falling and wage inflation rising when the value for the cointegrating equation increases (because of a combination of either higher headline inflation, lower wage inflation, or lower oil prices). Thus, the result clearly shows that price inflation (and oil prices) was closely linked to trend wage inflation developments before 1998. The coefficient on oil prices is also positive but it is not statistically significant.

We can gauge the impact of short-run wage movements on prices by looking at the coefficients on lagged ΔWAs in the equation for ΔPA. The table shows the estimates for both coefficients are negative, but they are not statistically significant and the magnitude is small compared with the size of coefficient on ΔWA in the cointegrating equation times the adjustment coefficient for price inflation. Thus, the short-run dynamics is not large enough to offset the long-run positive implication of wage inflation on price inflation.

Table 7 shows the estimates for the same system assuming two cointegrating relations over the same sample period (1981Q1-1997Q4). To make the interpretation easier, we impose the following restrictions on the cointegrating relations and the adjustment coefficients.

(i). The first cointegration equation does not include oil price and the coefficient on price inflation is normalized to be 1.

(ii). The second cointegration equation does not include wage inflation and the coefficient on the price inflation is normalized to be 1.

(iii). The adjustment coefficient for oil prices to the first cointegration equation is zero.

(iv). The adjustment coefficient on wage inflation to the second cointegration equation is zero.

Essentially these restrictions suggest that trend movements in headline prices are related separately to movements in oil prices and wages, and that oil prices and wages respond only to their own disequilibria vis a vis headline inflation. The two cointegrating equation estimated under these restrictions are:

$$\Delta PA - 1.4738*\Delta WA + 0.0087$$
$$(-5.90)$$

$$\Delta PA - 0.0024*PO + 0.0157$$
$$(-1.61)$$

The first one captures the long-run co-movements of wage inflation and price inflation, and the second one captures the long-run co-movement of oil prices and price inflation. The signs are expected: higher wages lead to higher prices and increases in the price of oil raise inflation, although the coefficient on oil price in the second equation is not statistically significant.

The signs of the adjustment coefficients in Table 7 all make sense: all the variables adjust as needed to re-establish the cointegrating relations. The coefficient for oil price is not significant, suggesting the adjustment to the second cointegration relation is mostly done by price inflation. Both price inflation and wage inflation seem to adjust to the first cointegrating relation, but the statistical significance of the coefficient for price inflation is only marginal.

The sum of the coefficients on lagged (changes in) wage inflation in the price equation is negative, but the magnitude is small again and they are statistically insignificant.

Taken together, the results in Tables 6 and 7 are consistent with the positive linkages between the trends in wage inflation and price inflation. The validity of these results, however, depends on the correct assumptions of the order of integration of the variables (PA and WA are I(2) and PO is I(1)) and the number of cointegration relations.

If we are just interested in Granger causality between these variables, Toda and Yamamoto (1995) provide a test that does not require knowledge of either the true degree of integration or the nature of cointegration. Giles (2011) is a very helpful exposition of this approach. The procedure amounts to estimating the VAR system in levels of the variables with *d* extra lagged endogenous variables where *d* is the maximum order of integration for the endogenous variables. Then, the standard Wald statistic on the coefficients excluding those on the extra *d* lags follows usual chi-square distribution asymptotically.

We apply this approach to the system of (PA, WA, PO) (note all of these are in levels) for 1981Q1-1997Q4 sample. First, we estimate unconstrained VAR with centered quarterly dummies, a linear trend and two VAT dummies (Tax89 and Tax97). The AIC suggests the optimal lag length is 11. Since the endogenous variables can be I(2), Toda-Yamamoto procedure requires adding at least two more lags to estimate the VAR and testing the exclusion of the first 11 lagged variables.

Table 8 shows the results of such a Granger causality test. At 5% significance level, we find the wages Granger cause prices, although prices do not cause wages. We also find the oil price Granger causes the headline CPI. Thus, Toda and Yamamoto approach also confirms that wage

levels influenced price levels during the 1981-1997 period. Although we do not report the coefficient estimates of the VAR, we note that two VAT dummies are statistically significant in the price equation while the trend variable does not come in significantly in any equation.

Now we repeat this analysis for the 1998-2017 period. Because the ADF tests in Table 1 suggest that headline CPI, average wage, and oil price in Japanese yen are all I(1) during this sample period, we consider a system of (PA, WA, PO). First, we test for cointegration. The results for the Johansen cointegration tests are reported in Table 9. Again the two statistics give us conflicting results. The trace test suggests these variables are actually I(0), which is inconsistent with the results of the Augmented Dickey Fuller tests. The maximum eigenvalue test finds one cointegrating vector.

We proceed by estimating an error correction model assuming 1 cointegration relation. The results are reported in Table 10. The estimated cointegration equation is:

$$\underset{}{PA} - \underset{(-7.07)}{1.3498}*WA + \underset{(8.15)}{0.0194}*PO$$

The positive and significant coefficient on oil price is puzzling since it implies that when oil prices rise, head line inflation would be expected to fall. The adjustment coefficient for headline CPI in Table 10 is also puzzling. The coefficient implies, for instance, if wages and prices are misaligned, then prices move farther apart (although the coefficient estimated is not statistically significant). The adjustment of wage level is in the right direction and the coefficient is statistically significant.

The short-run impact of wage inflation on price inflation makes sense. The point estimate of the coefficient on one quarter lag of wage inflation in the price inflation equation suggests that a one percentage point increase in wage inflation raises price inflation in the next quarter by 0.18%, although the statistical significance is marginal. Interestingly, this is much larger than the impact of the own lag of price inflation. The coefficient is actually negative, which is also puzzling.

When we apply Toda-Yamamoto approach to the system of (PA, WA, PO) in the 1998-2017 period, AIC suggests that the optimal lag length is 2. We again include centered quarterly dummies, a linear trend, and a VAT increase dummy (Tax14). Since the variables are likely to be I(1), we add one extra lag to the system and test for exclusivity of the first two lags. The result is reported in Table 11. At usual levels of significant, none of variables significantly causes any of the others. The VAT dummy (coefficient estimate not reported) is again significant in the price equation. Interestingly, the trend variable comes in significantly negative for the price level.

Overall, there are a number of puzzling results about the connections between wages and prices after 1998. Once the trend in wage inflation disappeared, the long-run link between wages and prices vanished too. Movements in wage inflation also ceased to predict movements in price

inflation. The link between headline price inflation and oil prices also broke. In short, the close and intuitive links between price and wage inflation that existed in the 1981-1997 period collapsed after 1998.

6. Conclusion

This paper looks at why Japan's inflation has not changed much despite the extremely aggressive monetary easing and the expansion of the real economic activity under Abenomics. Our explanation focuses on the role of dual labor market that made wages less sensitive to the labor market conditions. Japan's labor market has become increasingly segmented into full-time workers, many of whom are on lifetime employment and hence rarely fired, and part-time workers, whose employment can be adjusted more flexibly over business cycles. Following the recession in the late 1990s, many firms ended up hoarding an excessive number of full-time workers. Attempts by many firms to increase the proportion of part-time workers to gain the ability to adjust their employment more effectively in globally competitive markets, which preceded the recession, intensified the surplus of regular full-time workers.[4] The surplus of regular workers meant that there was no pressure on their wages. The wages for part-time workers do depend on business conditions, but since the level of their wages is much lower than full-time workers, the substitution of part-time for full-time employees has meant that the average wage for the economy has stagnated even as economic conditions have improved.

While these broad trends are known, especially for labor economists, their role in understanding inflation has been less studied. We provide some evidence on aspects of the story that bear on inflation. We find that the link between labor market conditions and wages that existed before 1998 was substantially weakened after 1998. Looking at wages for full-time workers and part-time workers separately in the period after 1998, we find the wages for full-time workers do not respond to tightness of the labor market while the wages for part-time workers are quite sensitive.

Thus, one reason why Japanese inflation does not seem to have responded to economic recovery is the decoupling of average wages and labor market conditions. Looking at the relationship between wages and prices, however, shows that this is not the whole story. We find that the link between wages and prices also changed around 1998. Before 1998, wage inflation was an important determinant of price inflation. The link, however, seems to have disappeared after 1998. Further understanding the causes of the disconnect between wages and prices after 1998 is an important topic for future research.

[4] For example, see Abe and Hoshi (2007) for some evidence on increased adjustment of employment and reduced emphasis on lifetime employment.

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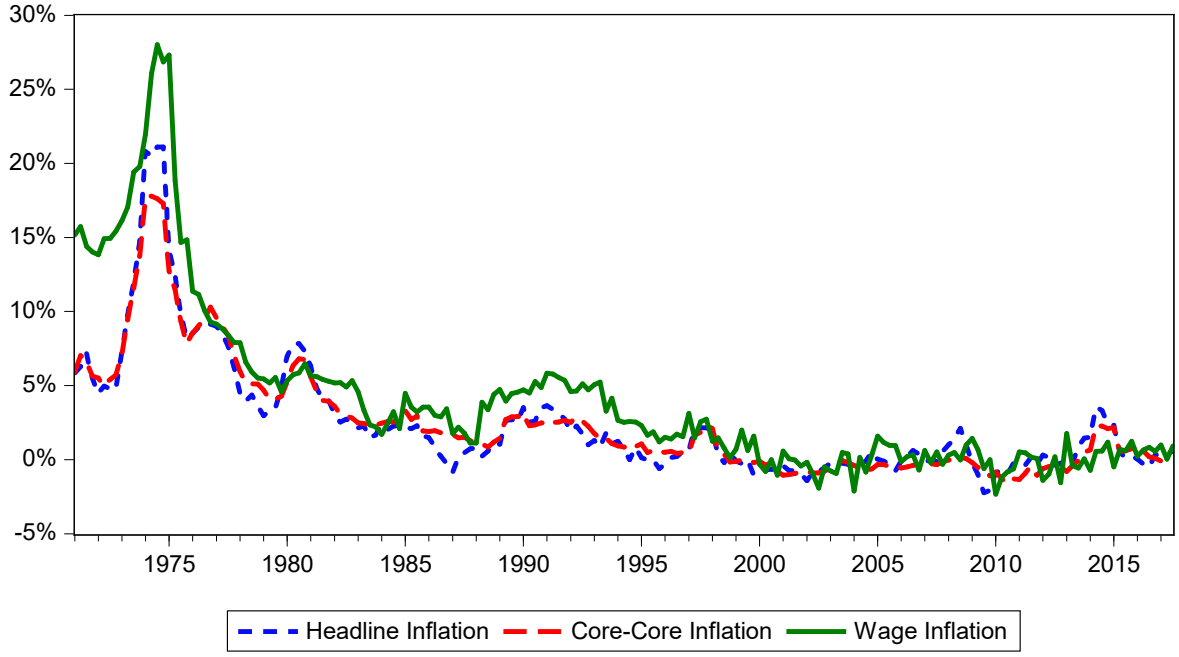
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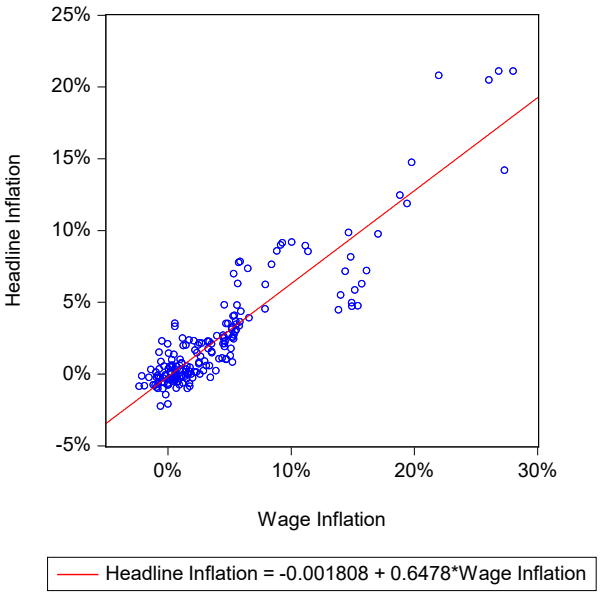
Figure 1: Japanese Price and Wage Inflation



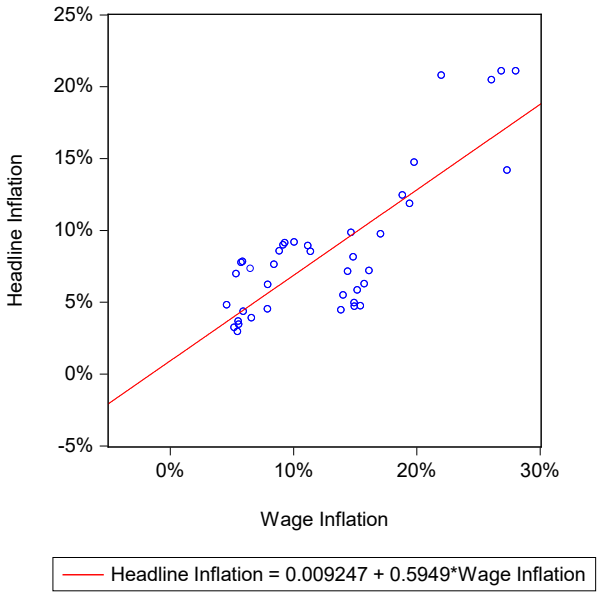
Source: Authors' calculation based on Ministry of Internal Affairs and Communications, Statistics Bureau, *Consumer Price Index*, and Ministry of Health, Labour and Welfare, *Monthly Labour Survey*.

Figure 2: Japanese Wage and Price Inflation over Different Periods

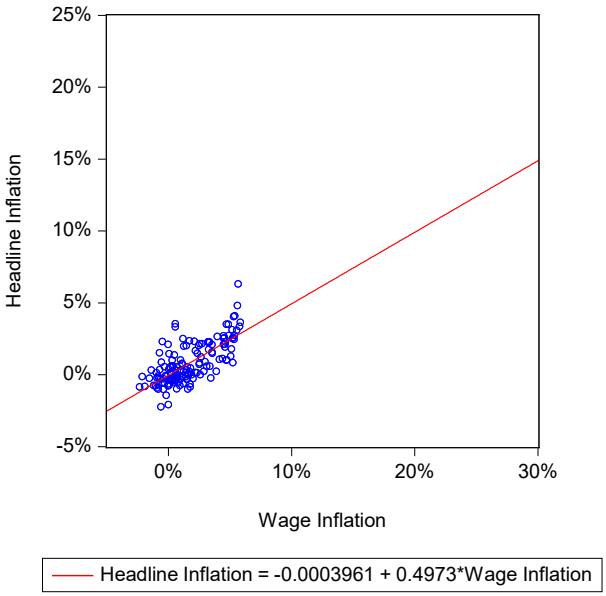
Panel A. 1971 - 2017



Panel B. 1971-1980

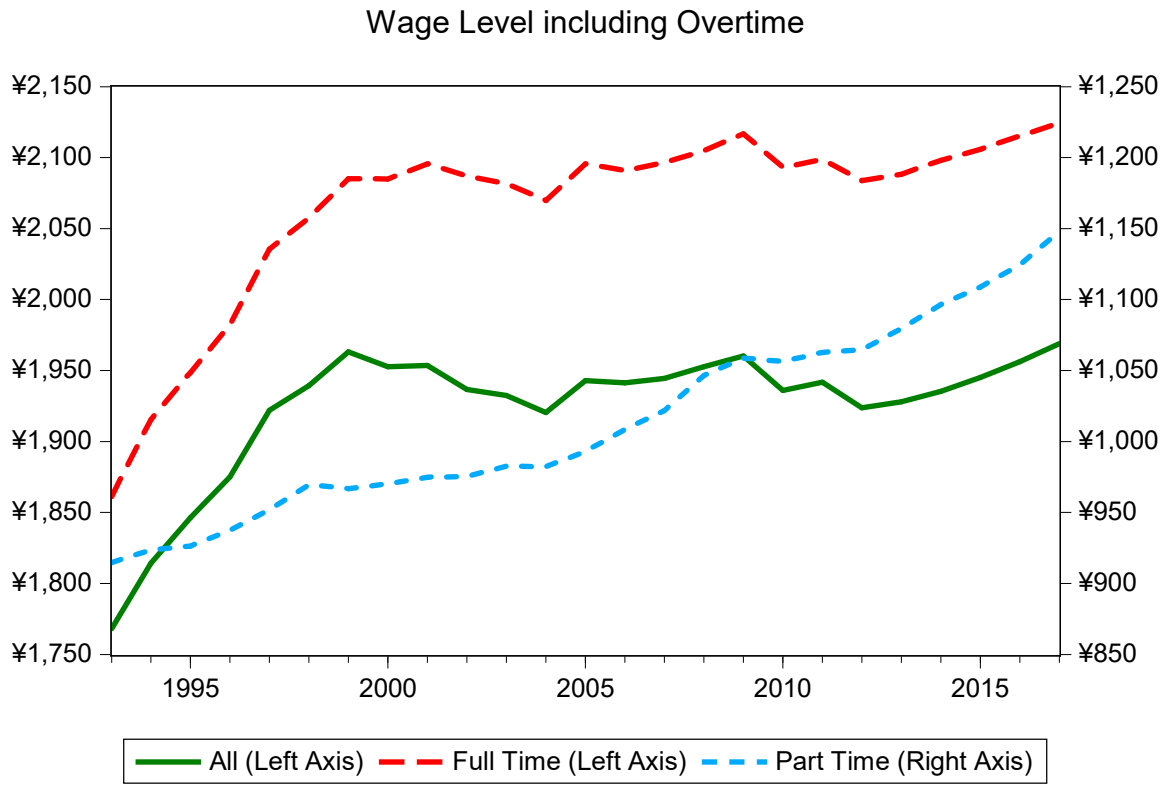


Panel C. 1981-2017



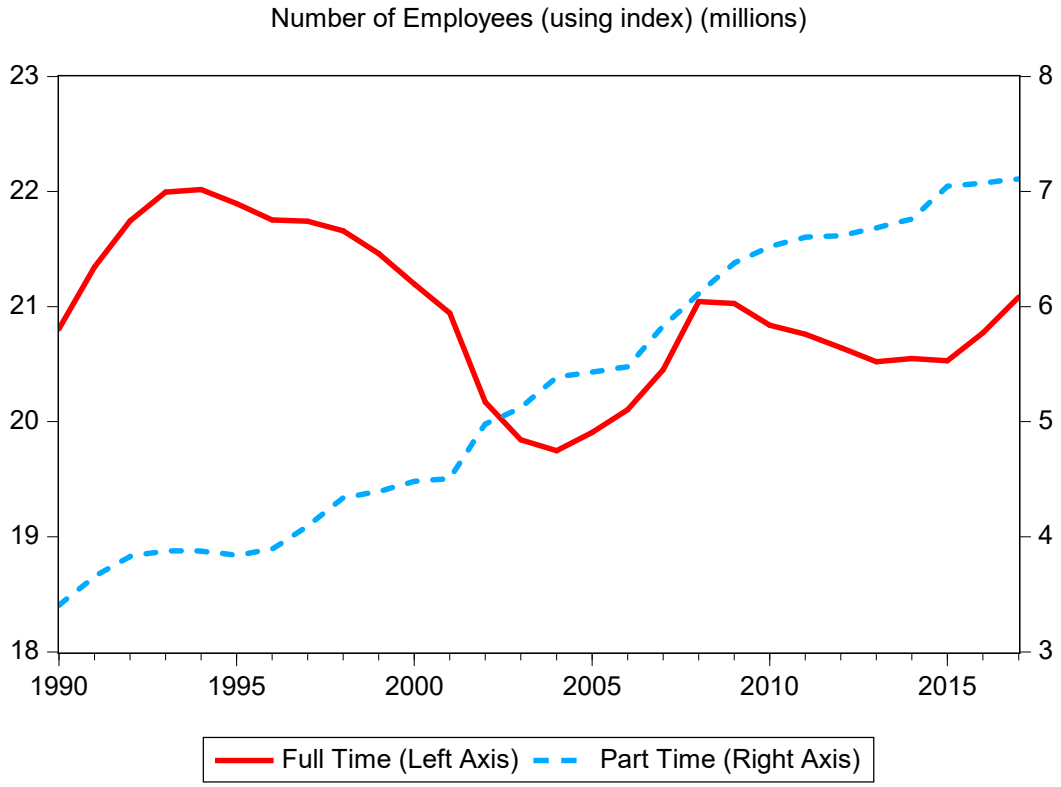
Source: Authors' calculation based on Ministry of Internal Affairs and Communications, Statistics Bureau, *Consumer Price Index*, and Ministry of Health, Labour and Welfare, *Monthly Labour Survey*.

Figure 3: Wages in Japan, 1993-2017



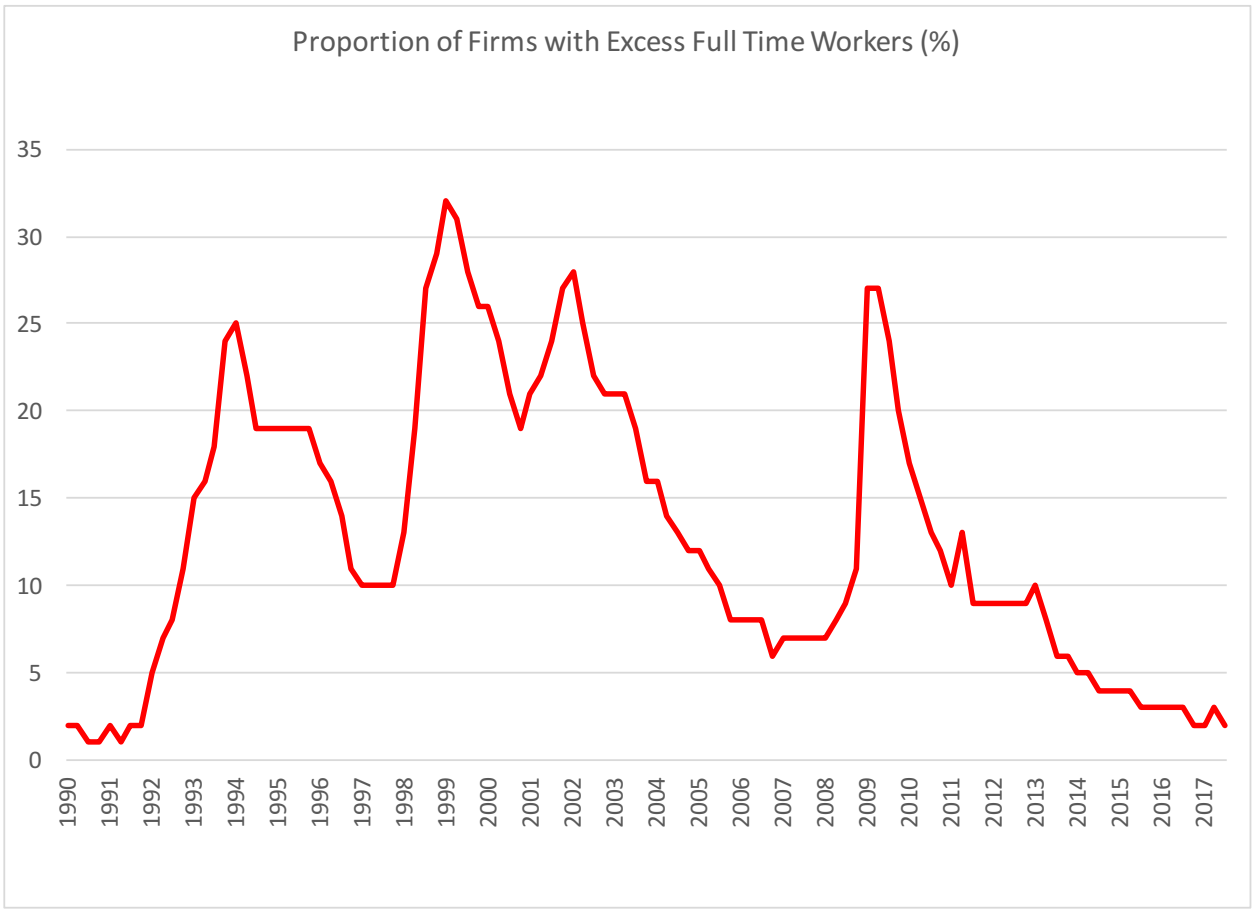
Source: Authors' calculation based on Ministry of Health, Labour and Welfare, *Monthly Labour Survey*.

Figure 4: Composition of the Japanese Workforce (Establishments with more than 30 employees)



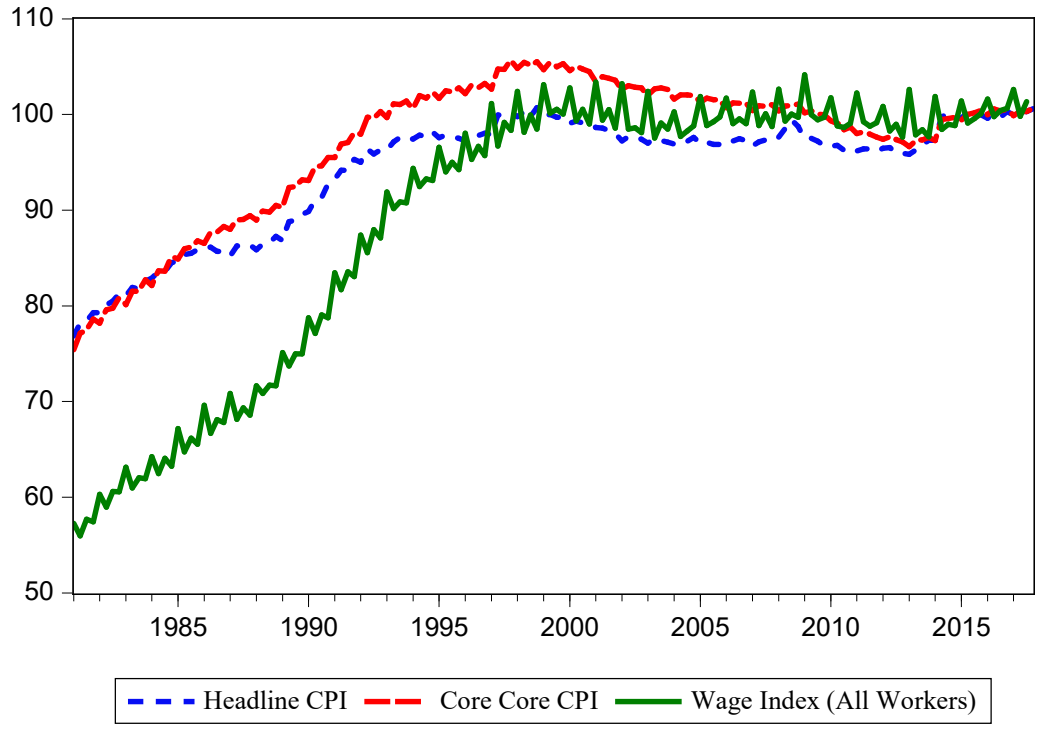
Source: Authors' calculation based on Ministry of Health, Labour and Welfare, *Monthly Labour Survey*.

Figure 5. Proportion of Firms with Excess Full Time Workers (%)



Source: Ministry of Health, Labour and Welfare, *Survey of Labour Economy Trend.*

Figure 6: Japanese Price and Wage Levels 1981-2017



Source: Ministry of Internal Affairs and Communications, Statistics Bureau, *Consumer Price Index*, and Ministry of Health, Labour and Welfare, *Monthly Labour Survey*.

Table 1: Augmented Dickey Fuller Tests for Orders of Integration

	Order of Integration	
Series	1981-1997	1998-2017
Log of Headline Consumer Price Index	I(2)	I(1)
Log of Core-Core Consumer Price Index	I(1)	I(1)
Log of Average Wages for All Workers	I(2)	I(1)
Log of Oil Price in Japanese Yen	I(1)	I(1)

Notes: The oil price is for West Texas Intermediary Crude Oil converted to yen using the quarterly average yen-dollar exchange rates. Both the oil price and exchange rate series come from FRED (Federal Reserve Economic Data) at the Federal Reserve Bank of St. Louis.

Table 2: ARMAX Models for Price and Wage Inflation

Headline CPI Inflation

	Sample: 1981 - 1997			Sample 1998- 2017		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	0.0064	0.0010	6.35	0.0003	0.0008	0.35
Q1 Dummy	-0.0069	0.0010	-7.13	-0.0041	0.0012	-3.47
Q2 Dummy	0.0026	0.0009	2.70	0.0032	0.0014	2.24
Q3 Dummy	-0.0063	0.0008	-7.92	-0.0010	0.0011	-0.90
1989Q2 Dummy	0.0185	0.0043	4.34			
1997Q2 Dummy	0.0139	0.0047	2.94			
2014Q2 Dummy				0.0210	0.0369	0.57
MA(2)	0.3591	0.1172	3.06	0.0695	0.1122	0.62
MA(3)	0.3323	0.1180	2.82	0.0493	0.1507	0.33
MA(4)	0.0958	0.1165	0.82	0.0518	0.1496	0.35
MA(5)	0.6753	0.1723	3.92	-0.1659	0.1305	-1.27
R^2	0.76			0.55		
Approximate P-value Q statistic	0.406 (at lag 17)			0.340 (at lag 19)		

Core-Core Inflation

	Sample: 1981 - 1997			Sample 1998- 2017		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	0.0091	0.0009	10.70	0.0005	0.0006	0.95
Q1 Dummy	-0.0135	0.0014	-9.57	-0.0072	0.0009	-8.35
Q2 Dummy	0.0053	0.0008	6.84	0.0038	0.0009	4.45
Q3 Dummy	-0.0089	0.0019	-4.61	-0.0020	0.0009	-2.28
1989Q2 Dummy	0.0094	0.0066	1.43			
1997Q2 Dummy	0.0078	0.0062	1.25			
2014Q2 Dummy				0.0166	0.006	2.70
MA(2)	0.4808	0.1069	4.50	0.0888	0.1171	0.76
MA(4)	0.5418	0.1076	5.04	0.3463	0.1184	2.92
R^2	0.92			0.86		
Approximate P-value Q statistic	0.289 (at lag 17)			0.571 (at lag 19)		

Wage Inflation

	Sample: 1981 - 1997			Sample 1998- 2017		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	-0.0057	0.0038	-1.49	-0.0049	0.0025	-1.97
Q1 Dummy	0.0539	0.0053	10.27	0.0377	0.0043	8.73
Q2 Dummy	-0.0206	0.0031	-6.53	-0.0288	0.0022	-12.93
Q3 Dummy	0.0256	0.0050	5.11	0.0116	0.0048	2.42
1989Q2 Dummy	-0.0022	0.0165	-0.13			
1997Q2 Dummy	-0.0149	0.0363	-0.41			
2014Q2 Dummy				0.0037	0.0154	0.24
AR(1)	-0.1974	0.1524	-1.29	-0.7268	0.1497	-4.86
AR(2)	0.2187	0.1518	1.44	-0.1538	0.1293	-1.19
MA(4)	0.3538	0.1605	2.20	0.0688	0.1149	0.60
R^2	0.96			0.93		
Approximate P-value Q statistic	0.202 (at lag 17)			0.114 (at lag 19)		

Table 3: ARMAX Models for Wage Inflation with Labor Market Conditions

Average Wage Inflation for All Workers

A. Lagged Active Job Openings to Applicants Ratio (AJOAR) as the Labor Market Condition

	Sample: 1981 - 1997			Sample 1998- 2017		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	-0.0038	0.0030	-1.26	-0.0042	0.0025	-1.65
Q1 Dummy	0.0538	0.0046	11.75	0.0375	0.0043	8.73
Q2 Dummy	-0.0217	0.0026	-8.25	-0.0287	0.0022	-13.10
Q3 Dummy	0.0256	0.0040	6.32	0.0115	0.0050	2.28
AR(1)	-0.4023	0.1615	-2.49	-0.7622	0.1430	-5.33
AR(2)	0.0423	0.1452	0.29	-0.1815	0.1318	-1.38
MA(4)	0.3105	0.1519	2.04	0.0518	0.1209	0.43
Lagged AJOAR	0.0083	0.0034	2.47	0.0026	0.0012	2.14
R^2	0.96			0.94		

B. Lagged Unemployment Rate (UNEMP) as the Labor Market Condition

	Sample: 1981 - 1997			Sample 1998- 2017		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	0.0137	0.0044	3.13	0.0044	0.0040	1.10
Q1 Dummy	0.0530	0.0043	12.43	0.0374	0.0042	8.81
Q2 Dummy	-0.0218	0.0025	-8.82	-0.0289	0.0020	-14.34
Q3 Dummy	0.0252	0.0038	6.64	0.0113	0.0050	2.24
AR(1)	-0.6087	0.1408	-4.32	-0.7808	0.1412	-5.53
AR(2)	-0.1477	0.1355	-1.09	-0.1975	0.1286	-1.54
MA(4)	0.2010	0.1358	1.48	0.0217	0.1161	0.19
Lagged UNEMP	-0.0202	0.0036	-5.66	-0.0062	0.0024	-2.62
R^2	0.97			0.93		

Table 4: ARMAX Models for Wage Inflation with Labor Market Conditions: Full Time and Part Time

Sample Period is 1998Q1-2017Q3

A. Full Time Wage Inflation

	Model 1			Model 2		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	-0.0051	0.0025	-2.02	0.0012	0.0041	0.30
Q1 Dummy	0.0405	0.0044	9.12	0.0404	0.0044	9.11
Q2 Dummy	-0.0306	0.0021	-14.39	-0.0307	0.0021	-14.94
Q3 Dummy	0.0138	0.0050	2.78	0.0137	0.0050	2.73
AR(1)	-0.7877	0.1391	-5.66	-0.8003	0.1357	-5.90
AR(2)	-0.1810	0.1271	-1.42	-0.1947	0.1254	-1.55
Lagged AJOAR	0.0015	0.0012	1.28			
Lagged UNEMP				-0.0045	0.0024	-1.91
R^2	0.94			0.94		

	Model 3			Model 4		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	-0.0049	0.0025	-1.93	-0.0056	0.0029	-1.90
Q1 Dummy	0.0405	0.0044	9.12	0.0405	0.0045	9.08
Q2 Dummy	-0.0306	0.0021	-14.44	-0.0306	0.0021	-14.67
Q3 Dummy	0.0138	0.0050	2.78	0.0138	0.0051	2.72
AR(1)	-0.7885	0.1390	-5.67	-0.7949	0.1357	-5.86
AR(2)	-0.1818	0.1270	-1.43	-0.1881	0.1284	-1.47
Lagged AJOAR for Full-Time Jobs	0.0014	0.0011	1.32	0.0008	0.0013	0.59
Lagged AJOAR for Part-Time Jobs				0.0019	0.0023	0.82
R^2	0.94			0.94		

B. Part Time Wage Inflation

	Model 1			Model 2		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	0.0028	0.0026	1.08	0.0144	0.0061	2.37
Q1 Dummy	0.0076	0.0047	1.62	0.0076	0.0053	1.44
Q2 Dummy	-0.0079	0.0015	-5.32	-0.0080	0.0018	-4.44
Q3 Dummy	0.0035	0.0049	0.72	0.0034	0.0052	0.66
AR(1)	-0.7476	0.1449	-5.16	-0.6105	0.1971	-3.10
AR(2)	-0.0352	0.1687	-0.21	0.1517	0.2337	0.65
MA(3)	-0.0484	0.2031	-0.24	0.1377	0.1975	0.70
Lagged AJOAR	0.0051	0.0010	4.85			
Lagged UNEMP				-0.0089	0.0032	-2.81
R^2	0.70			0.66		

	Model 3			Model 4		
Variable	Coefficient	Standard Error	T statistic	Coefficient	Standard Error	T statistic
Constant	-0.0001	0.0037	-0.02	0.0035	0.0028	1.28
Q1 Dummy	0.0083	0.0060	1.37	0.0075	0.0047	1.61
Q2 Dummy	-0.0078	0.0016	-5.02	-0.0079	0.0015	-5.20
Q3 Dummy	0.0040	0.0056	0.72	0.0034	0.0049	0.70
AR(1)	-0.5615	0.1362	-4.12	-0.7499	0.1529	-4.91
AR(2)	0.2710	0.1347	2.01	-0.0493	0.1872	-0.26
MA(3)	0.2637	0.1576	1.67	-0.0653	0.2189	-0.30
Lagged AJOAR for Full-Time Jobs				0.0046	0.0010	4.47
Lagged AJOAR for Part-Time Jobs	0.0052	0.0032	1.64	-0.0002	0.0019	-0.11
R^2	0.65			0.70		

Table 5: Cointegration Test of (ΔPA, ΔWA, PO): 1981Q1-1997Q4

# of cointegration vectors	Trace Statistic	Maximum Eigenvalue Statistic
None	48.02453 (0.0002)	31.60773 (0.0012)
At most 1	16.41680 (0.0362)	13.17352 (0.0738)
At most 2	3.243283 (0.0717)	3.243283 (0.0717)

Trace test indicates 2 cointegrating relations at the 0.05 level.
Maximum eigenvalue test indicates 1 cointegrating relation at the 0.05 level.

Notes: The sample period is from 1981Q1 to 1997Q4. The endogenous variables are assumed to have linear deterministic trends and the cointegrating equations are assumed to have intercepts. The number of lags of the test VAR is 2. Centered seasonal dummies and the first differences of Tax89 and Tax97 are included as exogenous variables. The number in parentheses below test statistic is the p-value according to MacKinnon, Haug, and Michelis (1999).

Table 6: Error Correction Model of (ΔPA, ΔWA, PO) with One Cointegrating Relationship: 1981Q1-1997Q4

Dependent Variable	Δ^2PA	Δ^2WA	ΔPO
CointEq1	-0.4936	1.2177	2.8792
	(-2.75)	(3.61)	(0.42)
Δ^2PA(-1)	-0.3852	-0.6177	-2.2511
	(-2.51)	(-2.15)	(-0.39)
Δ^2PA(-2)	-0.2456	-0.2721	-3.2246
	(-2.26)	(-1.33)	(-0.78)
Δ^2WA(-1)	-0.1940	-0.2668	0.0005
	(-1.56)	(-1.14)	(0.0001)
Δ^2WA(-2)	-0.0088	-0.0763	-3.3479
	(-0.11)	(-0.52)	(-1.14)
ΔPO(-1)	0.0103	0.0090	0.2607
	(2.98)	(1.39)	(2.00)
ΔPO(-2)	0.0064	-0.0060	-0.0579
	(1.67)	(-0.84)	(-0.40)
Δ^2TAX89	0.0077	0.0066	0.1207
	(2.59)	(1.18)	(1.07)
Δ^2TAX97	0.0042	-0.0036	-0.0572
	(1.41)	(-0.65)	(-0.51)
R-squared	0.9103	0.9876	0.1841

Notes: Sample period is 1981Q1-1997Q4. Each column shows the estimation result of the equation that explains the dependent variable in the first row by the variables specified in the first column. Each cell shows the coefficient estimate and its t-statistic on the row variable in the column equation. Each equation also includes centered quarterly dummies and a constant. CointEq1 is the cointegrating equation, estimated to be ΔPA – 0.8054*ΔWA – 0.0017*PO + 0.0165.

Table 7: Error Correction Model of (ΔPA, ΔWA, PO) with Two Cointegrating Relationships: 1981Q1-1997Q4

Dependent Variable	Δ^2PA	Δ^2WA	ΔPO
CointEq1	-0.1783	0.7851	0.0000
	(-1.76)	(4.06)	(NA)
CointEq2	-0.5126	0.0000	2.6878
	(-3.48)	(NA)	(0.46)
Δ^2PA(-1)	-0.2999	-0.4299	-2.0493
	(-2.01)	(-1.50)	(-0.35)
Δ^2PA(-2)	-0.2246	-0.2045	-3.0379
	(-2.16)	(-1.02)	(-0.74)
Δ^2WA(-1)	-0.1064	-0.1480	-1.4386
	(-0.86)	(-0.62)	(-0.29)
Δ^2WA(-2)	0.0250	-0.0324	-3.9479
	(0.33)	(-0.22)	(-1.32)
ΔPO(-1)	0.0095	0.0075	0.2662
	(2.86)	(1.17)	(2.01)
ΔPO(-2)	0.0072	-0.0051	-0.0600
	(1.97)	(-0.73)	(-0.41)
Δ^2TAX89	0.0068	0.0054	0.1261
	(2.36)	(0.97)	(1.10)
Δ^2TAX97	0.0036	-0.0047	-0.0581
	(1.25)	(-0.84)	(-0.51)
R-squared	0.9189	0.9883	0.1846

Notes: Sample period is 1981Q1-1997Q4. Each column shows the estimation result of the equation that explains the dependent variable in the first row by the variables specified in the first column. Each cell shows the coefficient estimate and its t-statistic on the row variable in the column equation. Each equation also includes centered quarterly dummies and a constant. CointEq1 is the first cointegrating equation, estimated to be ΔPA – 1.4738*ΔWA + 0.0087. CointEq2 is the second cointegrating equation, estimated to be ΔPA – 0.0024*PO + 0.0157.

Table 8: Granger Causality Tests of (PA, WA, PO) using the Toda and Yamamoto (1995) Approach: 1981Q1-1997Q4

Hypothesis	Wald Statitic	p-value
WA causes PA	20.20922	0.0426
PO causes PA	20.24071	0.0421
PA causes WA	14.15210	0.2247
PO causes WA	13.24760	0.2775
PA causes PO	10.34922	0.4993
WA causes PO	6.286510	0.8536

Notes: Sample period is from 1981Q1 to 1997Q4. Test VAR includes 13 lags: 11 lags suggested by AIC and 2 extra, which corresponds to the maximum order of integration for the variables in the system. The asymptotic distribution of each Wald statistic in the table is χ^2 with 11 degrees of freedom.

Table 9: Cointegration Test of (PA, WA, PO): 1998Q1-2017Q3

# of cointegration vectors	Trace Statistic	Maximum Eigenvalue Statistic
None	48.08711 (0.0002)	29.73890 (0.0024)
At most 1	18.34822 (0.0181)	13.04445 (0.0772)
At most 2	5.303768 (0.0213)	5.303768 (0.0213)

Trace test indicates 3 cointegrating relationships at the 0.05 level.
Maximum eigenvalue test indicates 1 cointegrating relation at the 0.05 level.

Notes: The sample period is from 1998Q1 to 2017Q3. The endogenous variables are assumed to have linear deterministic trends and the cointegrating equations are assumed to have intercepts. The number of lags of the test VAR is 2. Centered seasonal dummies and Tax14 are included as exogenous variables. The number in parentheses below test statistic is the p-value according to MacKinnon, Haug, and Michelis (1999).

Table 10: Error Correction Model of (PA, WA, PO) with One Cointegrating Relationship: 1998Q1-2017Q3

Dependent Variable	Δ^2PA	Δ^2WA	ΔPO
CointEq1	0.1143	0.3493	-5.2524
	(1.47)	(3.01)	(-1.84)
ΔPA(-1)	-0.1634	-0.3258	-0.4618
	(-0.94)	(-1.26)	(-0.07)
ΔPA(-2)	-0.0939	-0.1192	-3.4950
	(-0.58)	(-0.49)	(-0.59)
ΔWA(-1)	0.1759	-0.4416	-6.6979
	(1.82)	(-3.06)	(-1.89)
ΔWA(-2)	0.0513	-0.0249	-2.0351
	(0.70)	(-0.23)	(-0.76)
ΔPO(-1)	0.0081	-0.0087	0.1008
	(2.38)	(-1.71)	(0.81)
ΔPO(-2)	0.0029	-0.0004	-0.0725
	(0.80)	(-0.07)	(-0.55)
ΔTAX14	0.0008	-0.0057	0.0712
	(0.45)	(-2.03)	(1.04)
R-squared	0.4556	0.9502	0.3075

Notes: Sample period is 1998Q1-2017Q3. Each column shows the estimation result of the equation that explains the dependent variable in the first row by the variables specified in the first column. Each cell shows the coefficient estimate and its t-statistic on the row variable in the column equation. Each equation also includes centered quarterly dummies and a constant. CointEq1 is the cointegrating equation, estimated to be PA – 1.3498*WA + 0.0194*PO.

Table 11: Granger Causality Tests of (PA, WA, PO) using the Toda and Yamamoto (1995) Approach: 1998Q1-2017Q3

Hypothesis	Wald Statistic	p-value
WA causes PA	3.567903	0.1680
PO causes PA	4.406404	0.1104
PA causes WA	3.249765	0.1969
PO causes WA	3.498624	0.1739
PA causes PO	0.774479	0.6789
WA causes PO	0.703867	0.7033

Notes: Sample period is from 1998Q1 to 2017Q3. Test VAR includes 3 lags: 2 lags suggested by AIC and 1 extra, which corresponds to the maximum order of integration for the variables in the system. The asymptotic distribution of each Wald statistic in the table is χ^2 with 2 degrees of freedom.