Historical Wage Phillips Curves*

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Preliminary version: September, 2020

Abstract

Over the last years, many have been questioning the importance of the Phillips curve arguing that it has "flatten out of favour". Thus, a lot of attention has been given to understand why its slope is flatter and how can central bankers still explore it. Building on this current debate, I estimate historical wage Phillips curves by using newly assembled data on wages and unemployment rates for a set of 17 advanced economies starting in 1870. I show that the wage Phillips curve has always been "alive and well" but, similarly to recent times, it was flatter during the Gold Standard. The collected data suggest that a low price inflation environment promotes this disconnect, which is aligned with the New Keynesian model predictions. In such an environment, wages and prices are adjusted by firms less often and thus, the relationship between unemployment and wage inflation becomes weaker.

JEL classification: E01, E31, E52, N1, N2.

Keywords: Phillips curve, Phillips multiplier, wage inflation, unemployment, monetary policy, economic history

*I am indebted to Ana Sofia Pessoa, Andreas Westermark, Christian Bayer, David Vestin, Donghai Zhang, Evi Pappa, Felipe Valencia, Felix Ward, Francisco Amaral, Jean-Pierre Tabin, Joachim Jungherr, João Ritto, Julien Champagne, Kai Arvai, Leonardo Melosi, Mathias Klein, Maximilian Weiss, Moritz Schularick, Richard Hornbeck and Roberto Billi for their helpful comments and discussions. I likewise thank seminar participants at the University of Bonn. I would also like to thank the Sveriges Riksbank for hosting me while conducting part of the research that led to this paper. I acknowledge financial support from the FCT - Fundação para a Ciência e Tecnologia under the project Ref. SFRH/BD/144581/2019 and from the Deutsche Forschungsgemeinschaft (DFG, German Research Foundation) under Germany's Excellence Strategy - EXC 2126/1 - 390838866 and also under the RTG 2281 - The Macroeconomics of Inequality.

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A model in the Phillips curve tradition remains at the core of how most academic researchers and policymakers – including this one – think about fluctuations in inflation; indeed, alternative frameworks seem to lack solid economic foundations and empirical support. Donald Kohn (2008)

1 Introduction

The wage inflation-unemployment trade-off claims that changes in monetary policy push wage inflation and unemployment in opposite directions (Mankiw, 2001). Such relation is traditionally thought of in the form of a Phillips curve and is at the core of monetary policy. Most economists recognize its existence and central bankers rely on it to "transform" unemployment into inflation by making use of their policy interest rate (Barnichon and Mesters (2020b), Eser et al. (2020)).

Over the last years, many have been questioning the importance of both the wage and price Phillips curve arguing that they have "flatten out of favour". A flatter Phillips curve, as recently reported by numerous scholars (e.g. Del Negro et al. (2020)), could suggest that economic activity has a smaller effect on inflation and thus, central bankers' ability to steer inflation with policy-induced changes becomes smaller.

In this paper, I revisit the historical relationship between wage inflation and unemployment, the focus of Phillips (1958) original work, and make three main contributions.

First, I assemble historical annual data on nominal wages and unemployment rates starting in 1870 and covering a set of 17 advanced economies to answer these two questions. To the best of my knowledge, this is the first paper to bring such an historical perspective to the debate on the wage Phillips curve. Such approach keeps up with the recent trend of using long-run and cross-country perspectives to inform central debates in monetary and financial policy as in Reinhart and Rogoff (2009) and Schularick and Taylor (2012).

Second, I uncover considerable variation in the slope of the wage Phillips curve over time and find that its recent flattening is not a unique feature of the last 150 years. Third, I argue that changes in the price inflation environment is a possible cause of the observed time-varying slope. The data suggest that the relation between unemployment and wage inflation is flatter in times of low price inflation, which is aligned with the New Keynesian model predictions (Benati (2007)).

I start by reporting estimates of a micro-founded panel wage Phillips curve using exogenously identified monetary policy shocks as instruments in the spirit of Jorda and Nechio (2018). Here, I provide evidence that the wage Phillips curve has always been "alive and well" but, similarly to the last two decades, it was flatter during the Gold Standard. Adding to that, I find that in periods of low price inflation the relation becomes weaker. Such finding is robust across alternative specifications including different identification strategies. In my preferred estimates, a 1 percentage point (pp) increase in the monetary policy-induced unemployment rate is associated to a decrease in nominal wage growth of 0.4 pp. Such effect is reduced by almost 50% in a low price inflation environment.

The results carry on in a setting without the straitjacket of any assumed functional relation between wage inflation and unemployment. To be precise, I estimate a Phillips multiplier in the spirit of Barnichon and Mesters (2020b), similar to the impulse response-based statistic presented in Galí and Gambetti (2020). The main idea is to trace the evolution over time of the dynamic wage inflation-unemployment multiplier by looking at their impulse response functions to a monetary policy shock. While, on impact, the multiplier is about -0.3 pp, at longer horizons, the statistic becomes significantly more negative and divergent. Such large trade-off implies that a transitory policy-induced change in unemployment has a very persistent effect on wage inflation and thus, that central banks have sufficient ability to steer inflation with conventional monetary policy tools.

Still using the Phillips multiplier, I test the hypothesis that low price inflation weakens the wage inflation-unemployment trade-off via a state-dependent local projection instrumental variable (SDLP-IV) approach. The results show that, at a longer horizon of 3 years, the trade-off is smaller in periods of low price inflation. Thus, reinforcing the idea that in a low inflation environment, policy makers can not explore this trade-off to its full potential.

The New Keynesian model supports these findings. An increase (decrease) in trend inflation should cause an increase (decrease) in the frequency of price adjustment, and therefore a decrease (increase) in the steepness of the wage Phillips curve (Ball et al. (1988), Benati (2007)). This rationale in which low price inflation flattens the Phillips curve is in line with two other strands of the literature, namely the state dependent pricing (Costain, Nakov, and Petit (2019)) and the nominal price rigidities (Tobin (1972), Benigno and Ricci (2011), Daly and Hobijn (2014)).

Most of the recent empirical literature points out to a Phillips curve which is "alive and well", not only in the US (Coibion and Gorodnichenko (2015), Blanchard (2016), Höynck (2020)), but also in Europe (Levy (2019), Onorante et al. (2019)) and even worldwide (Coibion, Gorodnichenko, and Ulate (2019)).[1] Some also argue that it has becoming flatter in the recent years (Ball and Mazumder (2011), Galí and Gambetti (2020)).

Notwithstanding, since Ball et al. (1988), not enough attention has been given to the low

[1] A good summary of the literature since the inception of the Phillips curve can be found in Gordon (2011), while more recent discussions can be found in Mavroeidis, Plagborg-Møller, and Stock (2014) and Coibion, Gorodnichenko, and Kamdar (2018).

price inflation mechanism by the empirical literature. Some recent and notable exceptions are Benati (2007) who documents a positive correlation between the time-varying average gain of real activity and inflation; Vavra (2014) who rejects a New Keynesian Phillips curve with constant inflation output trade-off in favor of a slope that increases with microeconomic volatility; and Gertler and Hofmann (2018) who find a weaker money-inflation link in regimes characterized by low inflation.

This paper adds both a long-run and cross-country perspectives to the empirical literature on the existence of a time-varying wage Phillips curve by providing evidence that its slope varied significantly across time and that its flatness is associated with periods of low price inflation.

The remainder of this paper is organized as follows. Section 2 describes the data collection process and presents the descriptive statistics for the newly assembled series on wage inflation and unemployment rates. Section 3 discusses some of the most prominent identification challenges. Then, in section 4, I describe the empirical framework where I address the endogeneity concerns by controlling for changes in oil prices and by employing an instrumental variable approach. Here, I also run a formal test for changes in the slope of the wage Phillips curve in times of low price inflation. Section 5 addresses the specification uncertainty problem by performing a non-parametric estimation, the Phillips Multiplier, and provides further evidence that low price inflation is associated with a weaker wage inflation-unemployment trade-off. I conclude in section 6.

2 Data

I construct a new historical dataset composed of wage inflation and unemployment rates series that go as far as the nineteenth century in order to estimate historical wage Phillips curves. The newly assembled yearly data include a wage index measure and the unemployment rate for Australia, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Portugal, Spain, Sweden, the United Kingdom, and the United States. The sample spans 1870 to 2018 and Table 1 summarizes the data coverage by country.[2]

Table 1: Data Coverage

Country	Wages (nominal index)	Unemployment rate	CB Foundation
Australia	1870-2018	1901-2019	1911
Belgium	1870-2018	1921-2019	1850
Canada	1870-2018	1916-2019	1934
Denmark	1870-2018	1874-2019	1818
Finland	1870-2018	1920-2019	1811
France	1870-2018	1895-2019	1800
Germany	1870-2018	1887-2019	1876
Italy	1871-2018	1919-2019	1893
Japan	1870-2018	1930-2019	1882
Netherlands	1870-2018	1870-2019	1814
Norway	1870-2018	1904-2019	1816
Portugal	1870-2018	1953-2019	1846
Spain	1870-2018	1933-2019	1874
Sweden	1870-2018	1911-2019	1668
Switzerland	1870-2018	1913-2019	1907
United Kingdom	1870-2018	1870-2019	1694
United States	1870-2018	1890-2019	1913

Notes: This Table shows the earliest and the latest data point for each country's series. There are gaps in the unemployment rate data which mostly correspond to the war periods. All central bank foundations dates came from the central banks' websites.

The remaining macroeconomics data series used in this paper, such as the price inflation series, come from the Macrohistory Database (Jordà, Schularick, and Taylor (2017)). Given that before the Bretton Woods epoch only annual data is available for most variables, using panel data to study the wage Phillips curve is of paramount importance.

The wages and unemployment series can be broadly characterized as follows.

Unemployment rate: whenever possible, it is defined as the percentage of unemployed

[2]All data sources and further description of their construction are provided in the Online Appendix.

workers in the total labor force. According to Rasmussen and Pontusson (2018), most countries did not have any kind of unemployment insurance system until after the World Wars. Hence, citizens without a job had little incentives to enroll in a labor bureau given that there was no compulsory unemployment insurance.

The earlier data, which comes mainly from Mitchell (2013), Tabin and Togni (2013), Maddison (1982) and Galenson and Zellner (1957), build upon the previous caveat and present unemployment rates within smaller subsets of the active population such as trade unions or within people insured against unemployment. The assumption I am implicitly making when constructing a long-run series for unemployment is that the growth rate of the unemployment rates within smaller subsets of the active population is the same (or at least, highly correlated) to the unemployment rate in the country.

The most recent data follows the preferred definition and is based on either the Current Population Survey or the EU Labour force survey from the ILOSTAT. As a complement, data from the National Statistics agencies is used to ensure the robustness of this series.

Wages: whenever possible, wage series are an index of average earnings from all employees. However, the earlier data may build upon series for specific sectors according to the availability. I constructed this nominal index using old publications of statistical offices, financial history books and articles. The most recent data is based on the wage index series from the International Monetary Fund (IMF) and the Organization for Economic Cooperation and Development (OECD).

2.1 Descriptive statistics

Table 2 lists selected summary statistics of the dataset which was divided into five time windows and includes the number of observations, the mean, the standard deviation, the maximum and minimum of each series. The reasoning for such division is carefully explained in Appendix C. Both wage and price inflation series are computed as growth rates of a nominal index. The average wage inflation rate for the whole sample was 5.1%, almost 2 percentage points above the price inflation. The unemployment rate was, on average, 5.5%.

The lower means and standard deviations for wage and price inflation both during the first and last windows when compared to other periods stand out. All the studied countries were part of the Gold Standard agreement for most of the first time window as one can confirm in Table C.1. Even though only 7 out of the 17 countries in the sample are explicit inflation targeters (Svensson, 2010), it stems from Table 2 that using price inflation as the nominal anchor instead of the price of gold makes the volatility of price and wage inflation smaller

Table 2: Descriptive statistics

	N	Mean	Std. Dev.	Min	Max
1870-1913					
Unemployment rate	223	4.08	2.75	0.20	18.40
Wage inflation	223	1.68	2.63	-6.71	10.26
Price inflation	223	0.39	3.21	-10.94	11.56
1922-1938					
Unemployment rate	233	7.24	4.92	0.60	24.90
Wage inflation	233	1.35	5.83	-9.25	26.66
Price inflation	233	0.17	5.85	-15.32	30.43
1946-1971					
Unemployment rate	415	2.51	1.77	0.04	9.92
Wage inflation	415	7.75	5.02	-1.76	35.29
Price inflation	415	4.07	3.79	-6.87	20.38
1972-1994					
Unemployment rate	391	6.40	4.00	0.04	24.21
Wage inflation	391	9.23	6.23	-1.08	32.28
Price inflation	391	7.53	5.45	-0.71	37.88
1995-2018					
Unemployment rate	408	7.36	3.73	2.12	26.09
Wage inflation	408	2.51	1.68	-3.40	7.72
Price inflation	408	1.71	1.15	-1.35	5.24
Total					
Unemployment rate	1670	5.48	4.03	0.04	26.09
Wage inflation	1670	5.11	5.71	-9.25	35.29
Price inflation	1670	3.27	4.94	-15.32	37.88

Notes: All values are in percent. The war periods (1914-1921 and 1939-1945) are not included. This table only uses country-year observations for which there is data for the unemployment rate, and price and wage inflation. Table A.2 presents descriptive statistics for the unrestricted sample.

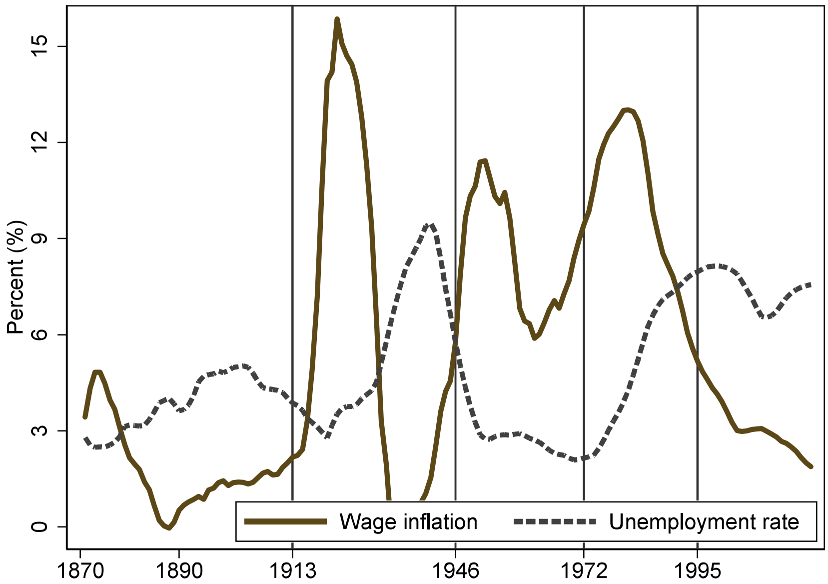
albeit the higher means.[3] Hence, the inflation targeting regime seems to be successfully keeping inflation under control with the lowest volatility ever observed.

Together with this Table, Figure 1 summarizes the cross-country trends in the data and points to a number of key insights which are relevant for the subsequent analysis. It plots a time-varying estimate of the mean wage inflation and the mean unemployment rate for the 17 countries in the sample using a 10-year rolling window. We can confirm that there is a strong negative comovement between the two variables.

During the Gold Standard, starting on the late 1890's we observe stable wage inflation and unemployment series. That picture dramatically changes once we enter in the war period

[3]The higher means should come without surprise given that targeting the price of gold implicitly yields a zero inflation expectation, contrarily to a 2% inflation target.

Figure 1: Mean wage inflation and unemployment rate



Notes: This figure plots a time-varying estimate of the mean wage inflation (solid olive line) and mean unemployment rate (dashed black line) using a 10-year rolling window using the full matched sample - when there is information for both variables.

with a large swing in the inflation series. The period between 1946 and 1971 corresponds to the Bretton Woods epoch and shows persistently low unemployment and high wage inflation rates. Then, after 1972, we can observe a peak for the inflation series which were partly driven by the two oil price shocks in 1973 and 1979. This peak is followed by a general decrease in inflation and an increase in unemployment stemming from the Great Moderation period.

Another way of looking at this correlation is to estimate a wage Phillips curve close to the original work of Phillips (1958) given by:

$$\pi^w_{c,t} = \alpha + \varphi u_{c,t} + \varepsilon_{c,t} \tag{1}$$

where $\pi^w_{c,t}$ is the annual wage inflation in country c at time t; α is a constant; $u_{c,t}$ is the unemployment rate in country c at time t; and $\varepsilon_{c,t}$ is an error term proxying for time-varying cost-push shocks to wages.[4] The twist of exploring the Phillips curve via a panel approach has been recently explored in Coibion et al. (2019), Levy (2019), and De Schryder, Peersman,

[4]The majority of the literature argues for the use of the unemployment gap instead of its level. However, that approach ignores the problem of measurement error arising from the computation of a natural unemployment rate. In my setting, due to the use of historical data, I believe that the latter poses a bigger threat because it is not possible to use detailed data to get the best estimates for the natural unemployment rate.

and Wauters (2020).

Figure 2: Wage Phillips curve

Notes: This figure plots a wage Phillips curve using annual data from 1870 to 2018 for all 17 countries in the main sample. It is a scatter plot with linear fit and a 95% confidence band. The dependent variable was truncated at the top 1% and bottom 1%.

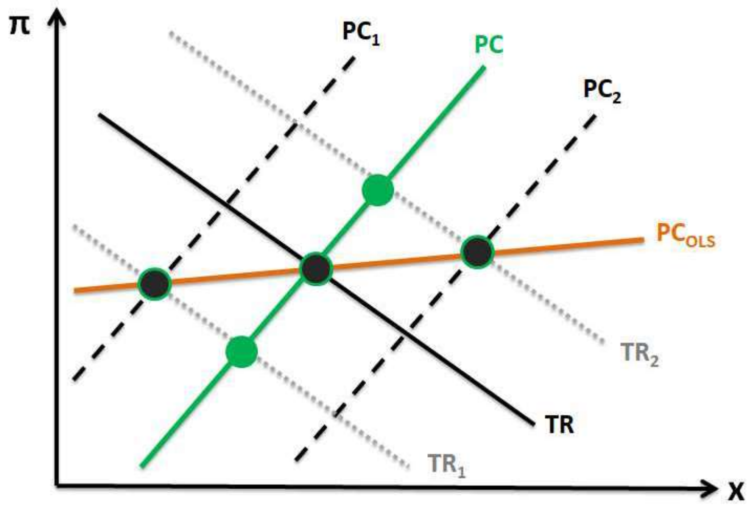
In Figure 2, we can see that wage inflation and the unemployment rate display a negative and significant correlation as predicted in the literature and in the previous analysis. Such finding can also be found at the country-level by looking at the last column of Table A.1.

3 Identification

The literature has extensively documented the empirical challenges in estimating both the price and wage Phillips curves (Mavroeidis et al. (2014) and McLeay and Tenreyro (2020)). The main concern is the **simultaneity bias** arising from the correlation between the measures of economic slack and inflation with the error term. Departing from an AS-AD model framework, there might be cost-push shocks affecting both the dependent and independent variables. These might be either shocks to inputs' prices - such as imported goods, oil or other important commodities - or inputs' quantities - such as a freeze in raw materials production or even wars which drain the labor force.

McLeay and Tenreyro (2020) make the case that the empirical disconnect between inflation and economic slack is expected to be emphasized when monetary policy is set optimally. To see this, Figure 3 displays a graphical relation between a measure of inflation on the y-axis and a measure of economic slack on the x-axis (for simplicity, take price inflation as π and output gap as x).

Figure 3: Phillips curve identification



Notes: Graphical representation of a targeting rule (TR) responding to cost-push shocks (1 and 2) which shift the true Phillips curve (PC). If one would infer a Phillips curve from observing equilibrium outcomes and using a simple OLS approach would obtain a flatter relationship (PC_{OLS}) between price inflation (π) and output gap (x).

On the one hand, the simultaneity bias is aggravated when central banks try to accommodate cost-push shocks. A negative cost-push shock (1) would shift the true price Phillips curve to the left ($PC \rightarrow PC_1$). A central bank would react by shifting its targeting rule in such a way that inflation becomes again close to its target ($TR \rightarrow TR_1$). The same analysis could be done to a positive cost-push shock (2). Overall, these observed equilibria would lead to the estimation of a flatter price Phillips curve (PC_{OLS}).

On the other hand, even absent of supply shocks, a purely inflation targeting central bank would neutralize any aggregate demand fluctuations to achieve constant inflation at its target. Hence, inducing a negative correlation between price inflation and economic slack and making it harder to uncover the true relationship between them McLeay and Tenreyro (2020).

The wage Phillips curve is less prone to these criticisms because many central banks do not explicitly target the unemployment rate. This observation is undeniably true for the majority of the sample in this study, only two central banks (from the USA and Australia) started targeting unemployment in the recent decades.

In the wage Phillips curve setting, both the unemployment rate and wage inflation might be correlated with the error term causing potential biases in the OLS estimates of a reduced form wage Phillips curve close to the one derived in Appendix B, Equation 17. Besides the cost-push shocks already mentioned, Galí (2011) and Galí and Gambetti (2020) argue that there could be wage markup shocks causing movements in both variables.[5]

In fact, the majority of the recent empirical literature estimating both price and wage Phillips curves agrees that identifying its slope requires exogenous shifts that cannot be explained by aggregate demand fluctuations. Acknowledging this issue, I control for important commodities price inflation as a proxy for cost-push shocks (Roberts (1995), Galí and Gertler (1999), Gali and Monacelli (2005)) and use an instrumental variable approach to isolate the remaining demand-driven variation in wage inflation.

I explore the *trilemma* of international finance by taking advantage of the fact that economies with fixed exchange rates are unable to implement independent monetary policies to identify exogenous changes to monetary policy as in Jorda and Nechio (2018).[6] However, when looking at such an historical sample, it might be harder to believe in "independent" central banks - Australia and the US did not even have a running central bank during the Gold Standard period. So, I borrow from Galí (2011), Galí (2015), and Coibion and Gorodnichenko (2015) and use lagged values of the economic slack variable (unemployment rate) as a second instrument to strengthen the identification strategy.

The second main challenge, which persists even after correcting for the simultaneity bias, is **specification uncertainty**. One can think of estimating a non-parametric version of the Phillips curve without the straitjacket of any ad-hoc functional relation between inflation and economic slack (Galí and Gambetti (2020)). Barnichon and Mesters (2020b),

[5]The natural wage markup is defined as the gap between the average real wage and the marginal rate of substitution that would prevail under flexible wages.

[6]Other possibility is using the current and lagged level of real government purchases of goods and services as an instrument (Roberts (1995) and Barnichon and Mesters (2020a)).

inspired by the fiscal multiplier literature (Ramey and Zubairy (2018)), propose estimating a Philips multiplier defined as the expected cumulative change in inflation caused by a demand shock that affects expected unemployment. This statistic directly captures the central bank's inflation-unemployment trade-off across different horizons, in line with the definition of Mankiw (2001). In section 5, I apply a similar methodology by making use of a Panel Local Projections Instrumental Variable (Panel LP-IV) approach (Jordà (2005)).

Currently, there is a large amount of **sampling uncertainty** with different researchers using different data vintages to compute Phillips curves. This work introduces two newly assembled historical data series on unemployment rates and wages for a set of 17 countries and a clean identification strategy in the hope of taking one step further to an empirical consensus.

Once again, the use of such a long-run panel is of utmost importance because it allows me to uncover whether the inflation environment is indeed an historical driver of the wage inflation-unemployment trade-off. Moreover, it also allows to explore more variation in wage inflation and thus, reduce the results' sensitivity to the data vintage that arises when using only one country and recent data.

4 Historical wage Phillips curves

In this section, I provide some reduced form evidence on the relationship between wage inflation and the unemployment rate in a panel of 17 countries. I tackle potential identification issues by controlling for important cost-push shocks and implementing an IV strategy making use of the *trilemma* instrument (Jordà, Schularick, and Taylor (2019)). Afterwards, I present rolling window estimates for the slope of the wage Phillips curve since the Gold Standard, highlighting its comovement with price inflation. Finally, I perform a formal parametric test to my hypothesis in the spirit of Coibion and Gorodnichenko (2015) and confirm that the slope of the wage Phillips curve is indeed flatter in periods of low price inflation.

I depart from the wage Phillips curve in equation (17) which was derived from the microfounded New Keynesian model in Appendix B and estimate:

$$\pi^w_{c,t} = \alpha + \varphi u_{c,t} + \gamma \pi^p_{c,t-1} + \mu_c + \epsilon_{c,t} \tag{2}$$

Relative to equation (1), I now introduce time-invariant country fixed effects μ_c to remove the role of differential local trends, and lagged price inflation, the measure to which wages are indexed. In accordance to the empirical literature, I am not including year fixed effects. Here, I am implicitly assuming that when there is no reoptimization, wages are indexed to ($\pi^p_{c,t-1}$) and γ represents the degree of indexation on past price inflation.[7]

Given an increase in the price level in $t-1$, workers bargain for a higher wage in t due to an increase in the cost of living in $t-1$. From Table A.1, we can see a correlation between price inflation in $t-1$ and wage inflation in t of more than 0.5 for almost all countries.

With this specification in mind, I now address the simulteneity bias problem. In summary, the identification challenge discussed in Section 3 arises from the presence of cost-push shocks to the Phillips curve, which is then magnified by the actions of monetary policymakers (McLeay and Tenreyro (2020)).

Controlling for supply shocks as suggested by Phillips (1958) and Gordon (1982) is one way to mitigate endogeneity problems. To do this, I use a readily identifiable source of $\epsilon_{c,t}$ variation by including the change in the oil price as a regressor to specification (2) as in

[7] Another possible interpretation is that firms look at the previous period price inflation as a good measure of inflation expectations which then affects their decision in changing both their products' prices and workers' wages.

Roberts (1995).[8] Thus, I estimate:

$$\pi^w_{c,t} = \alpha + \varphi u_{c,t} + \gamma \pi^p_{c,t-1} + c\pi^o_t + \mu_c + \epsilon_{c,t} \quad (3)$$

where π^o_t is the annual inflation of the international oil price in year t. The inclusion of oil prices can also account for inflation pressures coming from factors outside of each central bank's control. Moreover, oil prices are useful because they are determined in international markets and therefore, do not react to each country monetary policy (Jorda et al., 2019).[9] There is available data on the historical crude oil prices in US dollars per barrel from 1861 onwards in the Statistical Review of World Energy developed by BP.[10]

Adding to that, I employ an IV approach to address the endogeneity concerns and take it as my main specification. Countries with fixed exchange rates and open capital accounts are forced to track their base country interest rate movements. Hence, I use the base country interest rate movements as an instrument for the change in the short-run interest rate of each country following Di Giovanni, McCrary, and Von Wachter (2009), Jorda and Nechio (2018) and Jordà et al. (2019). This way, I am using an exogenous measure to the business cycle to instrument the unemployment rate. Such measure is straightforward to implement in my historical panel data setting. I take the instrument estimates directly from Jordà et al. (2019) and thus, I refer to their paper in regard to its construction. I also refer to the paper from Jorda and Nechio (2018) for a direct application of the instrument in the context of the Phillips curve estimation.

Notwithstanding, the *trilemma* instrument is not sufficiently strong to identify changes in the unemployment rate, especially during the Gold Standard epoch (F-Statistic is below 10). To solve this, I use lagged values of the unemployment rate as an extra instrument for its present value as in Galí (2011), Galí (2015), and Coibion and Gorodnichenko (2015).

Table 3 reports the estimates of φ, γ and c in equations (1), (2), and (3). I include the point-estimate of the coefficients, their robust standard errors clustered at the country level, and the R-squared of the estimation.

The last column reports the preferred specification, in which I address the simultaneity bias mentioned above using monetary policy shocks as instruments. It shows that we can predict an impact of around 0.38 percentage points lower wage growth per additional percentage point of the unemployment rate. It also shows that both oil and price inflation are relevant to explain wage inflation. This last piece of evidence allows me to support the idea

[8]The results go through when controlling for lagged oil price inflation. They also go through when using coal prices in the UK and US until the World Wars instead. However, given the high costs of transportation of the latter and the lack of data at the country level, I opt not to interpret these results.

[9]In Appendix, Figure A.1 presents the evolution of the international oil prices.

[10]You can find more information about this report here, consulted on 17^{th} July, 2019

Table 3: Wage Phillips curve estimates

Equation	(1)	(2-OLS)	(2-OLS)	(3-OLS)	(3-IV)
Unemployment (φ)	-0.515***	-0.308***	-0.354***	-0.333***	-0.379***
	(0.09)	(0.07)	(0.07)	(0.07)	(0.05)
Lagged Inflation (γ)		0.651***	0.633***	0.618***	0.652***
		(0.04)	(0.04)	(0.04)	(0.05)
Oil inflation (c)				0.030***	0.025***
				(0.00)	(0.00)
Country FE	NO	NO	YES	YES	YES
R^2			0.457	0.483	0.495
Observations	1811	1809	1809	1809	1321

Notes: Estimation of equations (1), (2), and (3) with robust standard errors in parentheses clustered at the country level. The dependent variable was truncated at the top 1% and bottom 1%.

that wages have some degree of indexation to lagged price inflation as argued in Appendix B.

I uncover a strong negative relationship between wage inflation and unemployment across all specifications. By taking this long-run perspective, I am able to take the focus out of the latest decades where Central Banks were active in policy making and truly assess the link between the real and nominal economies without only relying on an IV approach.

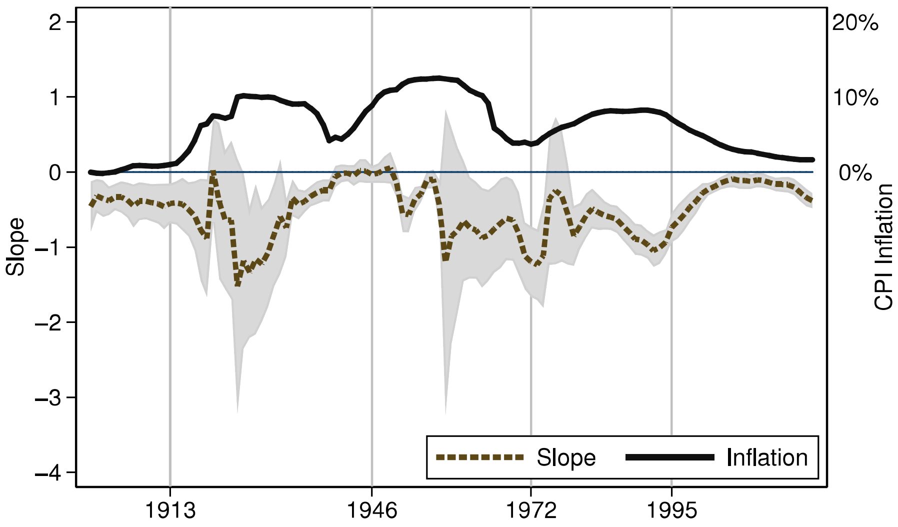
4.1 Rolling windows

In this subsection, I turn my attention to study the wage Phillips across historical periods. Figure 4 shows the time-varying estimates of its slope (φ) based on a Panel-IV regression with a 20-year rolling window. The estimates present evidence in favor of the hypothesis that in periods with lower price inflation, the slope of the wage Phillips curve is significantly flatter.

It is already documented, especially for the US, that the Phillips curve has been flattening (Ball and Mazumder (2011), Blanchard, Cerutti, and Summers (2015), Blanchard (2016), and Galí and Gambetti (2020)). What no one has ever documented before this paper is that the wage Phillips curve was also flatter during the Gold Standard period – a period with persistently low inflation.

Taking this finding at face value, the key takeaway is that today's Central Banks are not driving the flattening *per se*. Thus, arguing that Central Bank independence and the recent policy-making is driving the recent flattening seems to be valid as long as the underlying mechanism is based on their goal of targeting a low level of price inflation.

Figure 4: Panel-IV 20-year Rolling Window



Notes: This figure plots a time-varying estimate of the slope of the wage Phillips curve (φ) using annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on a rolling IV regression with a 20-year window and a 90% confidence band. The dependent variable was truncated at the top 1% and bottom 1%. Estimates for the Kleibergen-Paap Wald rk F-statistic in Figure A.4.

Figure 4 also displays the consecutive steepening after the end of the Bretton Woods agreement in 1972 and flattening of the wage Phillips curve afterward.[11]

Let us now take a look at the time-varying estimate of the lagged price inflation coefficient (γ) in Figure 5. It corroborates most of the literature arguing that the indexation degree of past inflation was lower or closer to zero during Gold Standard and the recent years (Benati (2008), Galí and Gambetti (2020)). This suggests that the importance of price indexation in wage setting is lower in lower inflation environments.

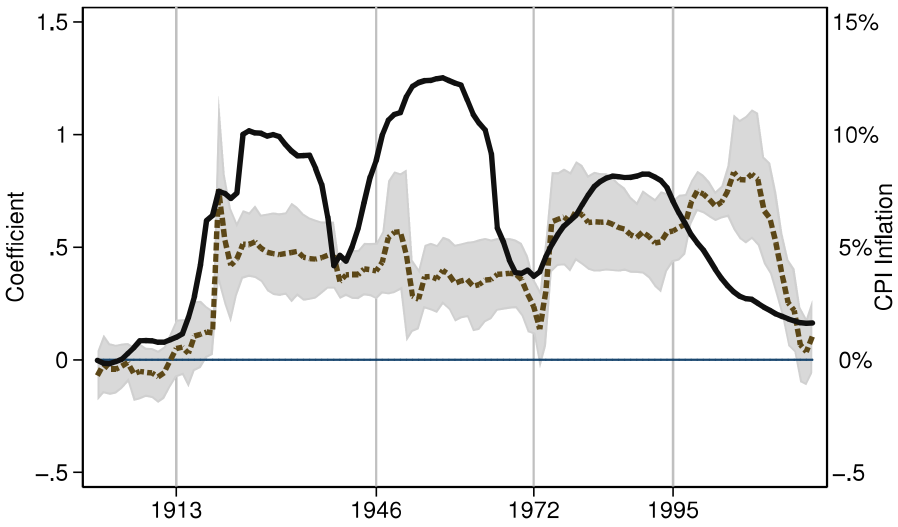
4.2 Low inflation flattens the Phillips curve

Several explanations have been put forward to explain the recent flatness in the Phillips curve, to name a few: globalization (Auer, Borio, and Filardo (2017)), central bank independence (Alesina & Summers, 1993), the role of China (Eickmeier and Kühnlenz (2018)), the financialization of commodity markets (Fernández, Schmitt-Grohé, and Uribe (2017)), firms' credit constraints (Gilchrist, Schoenle, Sim, and Zakrajšek (2017)), and the type of exchange rate regime (Geerolf (2020)).

In this paper, I test the low price inflation environment explanation that nests some of

[11]Such picture can be corroborated from the equivalent Panel-OLS estimate in Figure A.3. Albeit smaller in magnitude, the variation of the slope resembles the one presented in Figure 4.

Figure 5: IV 20-year Rolling Window



Notes: This figure plots a time-varying estimate of the lagged price inflation coefficient (γ) using annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on a rolling IV regression with a 20-year window and a 90% confidence band. The dependent variable was truncated at the top 1% and bottom 1%.

the above hypotheses – as the level of price inflation decreases, wages and prices are adjusted less often, leading to a smaller response of inflation to labor market conditions (Blanchard, 2016).[12]

To test whether low price inflation indeed affects the wage inflation-unemployment trade-off, I follow Coibion and Gorodnichenko (2015). I allow for a different slope of the wage Phillips curve when countries are in periods of low inflation as follows:

$$\pi^w_{c,t} = \alpha + \varphi \hat{u}_{c,t} + \Delta^{LI} \hat{u}_{c,t} I_{LI,t} + \delta^{LI} I_{LI,t} + \gamma E\{\pi^w_{c,t+1}\} + c\pi^o_t + \mu_c + error_{c,t} \tag{4}$$

where $I_{LI,t}$ is a dummy variable equal to one for periods when countries have price inflation below the threshold of 2% and above -2%, and 0 otherwise. The interaction of this dummy with the unemployment rate allows us to assess whether the slope of the Phillips curve is different in periods of low inflation.

The choice of this threshold is based on two facts: first, most central banks of advanced economies currently have a (close to) 2% inflation target; second, the median inflation in my sample is 2.2%. If I am to define what is a low inflation environment, irrespective of whether that changes throughout time, it has more relevance to look at periods that display

[12] In Appendix B.1, I try to shed more light on this argument departing from the New Keynesian model predictions.

inflation below 2%. Naturally, because I am implicitly assuming symmetry throughout my estimations, the lower bound should be -2%, or else I might be evaluating periods of strong deflation.[13]

Table 4 reports the estimation results of the wage Phillips curve both using a pooled OLS and an IV approach. Following the same rational as in the previous subsection, besides the *trilemma* instrument, I use as external instruments one lag of unemployment and the interaction of the dummy variable with the lag of the unemployment rate as in Coibion and Gorodnichenko (2015).

Table 4: Formal test for a break in the IV wage Phillips curve

	OLS	IV
Unemployment (φ)	-0.366***	-0.430***
	(0.08)	(0.07)
Interaction (Δ^{LI})	0.155**	0.242***
	(0.07)	(0.07)
Low Inflation (δ^{LI})	-2.721***	-3.219***
	(0.53)	(0.50)
Lagged Inflation (γ)	0.577***	0.608***
	(0.05)	(0.05)
Oil inflation (c)	0.027***	0.024***
	(0.00)	(0.00)
R^2	0.502	0.521
F-Statistic		670
Observations	1809	1321

Notes: Estimation of equation (4) by OLS and IV with robust standard errors in parentheses clustered at the country level. The dependent variable was truncated at the top 1% and bottom 1%. I provide the Kleibergen-Paap Wald rk F statistic which refers to the first-stage test for weak identification of excluded instruments.

The point estimates on the interaction term is always positive and significant hence, the wage Phillips curve slope was significantly and consistently flatter during low inflation periods. Therefore, the evidence for a change in the slope is clear and relatively large: on the order of a 50% reduction.

[13] I tried different specifications such as excluding all periods with inflation below 0%. I also tried adjusting the upper threshold between 1 and 3% and found that the lower the upper bound the more significant the results become. All in all, the qualitative findings go through. These results can be provide by the author.

5 Phillips Multiplier

In this section, I estimate the Phillips multiplier in the spirit of Barnichon and Mesters (2020b). It is a statistic that non-parametrically characterizes the trade-off between wage inflation and unemployment. This exercise allows me to tackle the problem of model mis-specification that might be present in the previous section. It also allows me to study the dynamics of wage inflation and unemployment responses to a policy-induced 1 pp increase in unemployment, i.e., a change orthogonal to the business cycle. Finally, I test the main hypothesis of the paper by exploring a state-dependent LP-IV estimation.

5.1 The multiplier

I start by providing inference on the Phillips multiplier based on an instrumental variable regression of cumulative inflation on cumulative unemployment using monetary policy shocks as the main instrument. I estimate an equation close to equation (4) in Barnichon and Mesters (2020b). The Phillips multiplier ($\mathcal{P}_h$) can be estimated using a Panel LP-IV approach from the following cumulative regression:

$$\sum_{j=0}^{h} \pi^{w}_{i,t+j} = \alpha_{i,h} + \mathcal{P}_h \sum_{j=0}^{h} \hat{u}_{i,t+j} + \zeta_h \mathbf{W}_{i,t} + \epsilon_{t+h} \tag{5}$$

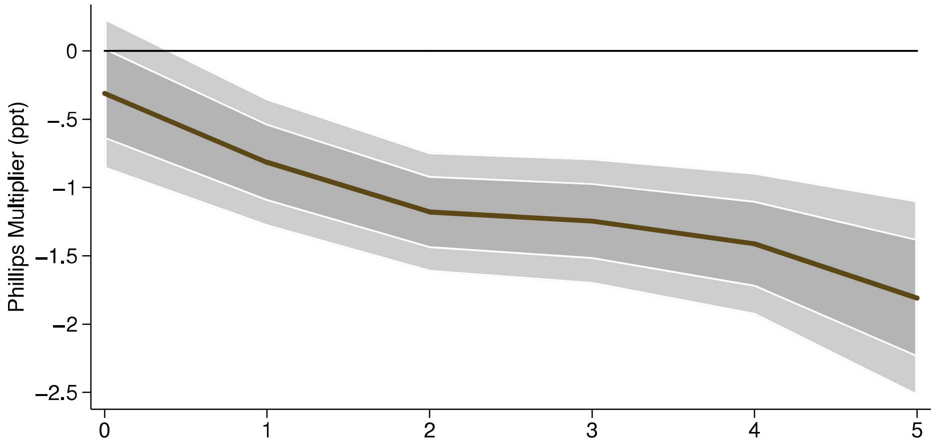
where $\alpha_{i,h}$ are country fixed-effects, and $\mathbf{W}_{i,t}$ is a vector of the control variables used in the previous section: oil price inflation and lagged price inflation, and also one lag of both wage inflation and unemployment rate as common in the estimation of local projections (Jordà (2005), Stock and Watson (2018)).

$\sum_{j=0}^{h} \hat{u}_{i,t+j}$ is instrumented by $\eta_{i,t}$, the exogenous changes in the short term interest rate obtained by using the *trilemma* instrumental variable first.[14] These monetary shock are orthogonal to supply shocks and to the natural unemployment rate under the common assumption that monetary policy is neutral under flexible prices (Galí, 2015). Via this IV approach, as in Barnichon and Mesters (2020b), the Phillips multiplier will allow me to estimate the trade-off (i) without bias from confounding supply shocks and (ii) without any need for a measure of the natural unemployment rate.

Figure 6 displays my estimate for the Phillips Multiplier for horizon $h = 0$ until $h = 5$. On impact, the multiplier is about -0.3 pp and significant at the 10% level, in line with what I found in the previous Section. Moreover, at longer horizons, the statistic becomes significantly more negative reaching about -1.25 pp after 3 years, diverging further on.

[14] I have argued before that this instrument is not strong enough for the first decades of the sample however, one should keep in mind that lagged unemployment is already being included as an instrument.

Figure 6: Phillips Multiplier



Notes: This figure plots the Phillips Multiplier estimated using the exogenous changes in the short term interest rate as instrument. The sample is annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on Equation 5 via a Panel LP-IV approach and displays a 68% and 90% confidence bands. The dependent variable was truncated at the top 1% and bottom 1%. Estimates for the Effective F statistic applying the procedure in Olea and Pflueger (2013) are in Figure A.5.

As Barnichon and Mesters (2020b) note, a large trade-off in the longer-run implies that a transitory policy-induced change in unemployment has a persistent effect on wage inflation. Hence, Figure 6 tells us that, throughout the analyzed sample, Central Banks had sufficient ability to steer inflation.

Figure A.5, in the appendix, reports the Olea and Pflueger (2013) F-statistics from the first-stage regression of equation 5 and documents that monetary policy shocks are correlated with cumulative unemployment at all horizons.[15] The F-statistic estimates are above the threshold of Olea and Pflueger (2013) for the first three periods and therefore, I am not relying on weak instrument robust methods for computing the confidence sets for the Phillips multiplier.

Finally, it is worth emphasizing the importance of the identification strategy. In Figure A.6, I present both the Phillips multiplier and its unconditional counterpart estimated using OLS. On impact, the unconditional multiplier is about -0.5 pp and significant at the 1% level, bigger than the conditional estimate and again in line with what I show in the previous section. However, contrary to the path of the Phillips multiplier, its counterpart displays no larger long-run trade-off reaching about -0.6 pp after 3 years, with no sign of divergence further on. Such differences caution against the use of OLS regressions to estimate the inflation-unemployment trade-off.

[15] I make use of the Stata command from Pflueger and Wang (2015) to compute the F-Statistics and the respective threshold.

5.2 Dynamics

In this subsection, I compute impulse response functions of the unemployment rate and wage inflation to a 1 pp increase in unemployment caused by exogenous changes in the monetary policy rate. To be precise, I estimate:

$$\sum_{j=0}^{h} X_{i,t+j} = \alpha_{i,h} + \beta_h^X \hat{u}_{i,t} + \zeta_h \mathbf{W}_{i,t} + \epsilon_{t+h} \tag{6}$$

where X is either the cumulative unemployment u or wage inflation π^w; $\beta_h^X \hat{u}_{i,t}$ is instrumented by $\eta_{i,t}$ the exogenous changes in the short term interest rate obtained by using the *trilemma* instrumental variable first, and the remaining terms are the same as in Equation 5.

Figure 7: Impulse Response Functions of Unemployment and Wage Inflation

Notes: This figure plots the impulse response functions of the unemployment rate and wage inflation to a 1 ppt increased in unemployment caused by exogenous changes in the short term interest rate. The sample is annual data from 1870 to 2018 for all 17 countries. It is computed based on Equation 6 via a Panel LP-IV approach and displays a 68% and 90% confidence bands. The dependent variable was truncated at the top 1% and bottom 1%. Effective F statistics applying the procedure in Olea and Pflueger (2013) are 42.54 which correspond to the impact F-Statistic in Figure A.5.

The key takeaways from Figure 7 are that unemployment response stops increasing after

$t = 3$ while wage inflation cumulative response is always increasing. Such behavior leads to a always decreasing multiplier.

These impulse response functions are estimated in such a way that we can obtain the Phillips Multiplier ($\mathcal{P}_h$) directly by dividing the unemployment coefficients from (6):

$$\mathcal{P}_h \equiv \frac{\beta_h^{\pi^w}}{\beta_h^u}$$

The advantage of doing the one-step estimation of the Phillips multiplier is to directly obtain the correct confidence bands, nevertheless the two-step estimation is consistent once samples are matched (Ramey and Zubairy (2018) and Barnichon and Mesters (2020b)).

5.3 Does low inflation also "flatten" the Phillips Multiplier?

Next, I re-do the test from subsection 4.2 by following closely Ramey and Zubairy (2018) whom provide a one step equation for the state-dependent LP-IV, which I adapt to:

$$\sum_{j=0}^{h} \pi_{i,t+j}^{w} = I_t \left[\mathcal{P}_h^L \sum_{j=0}^{h} \hat{u}_{i,t+j} + \zeta_h^L \mathbf{W}_{i,t} \right] + (1 - I_t) \left[\mathcal{P}_h^H \sum_{j=0}^{h} \hat{u}_{i,t+j} + \zeta_h^H \mathbf{W}_{i,t} \right] + \alpha_{i,h} + \epsilon_{t+h} \quad (7)$$

where I_t is an indicator for low inflation defined as a dummy variable equal to one for periods when countries have price inflation below the threshold of 2% and above -2%, and 0 otherwise. This exercise allows me to compare the evolution of the Phillips Multiplier in times of low versus high price changes (both inflation and deflation) and test directly whether $\mathcal{P}_h^L = \mathcal{P}_h^H$.

Table 5: State Dependent Phillips Multipliers

	Impact	1-Year	2-Years	3-Years
$\mathcal{P}_h^L$	-0.44	-0.51	-0.49	-0.55
	(0.35)	(0.39)	(0.38)	(0.35)
$\mathcal{P}_h^H$	-0.22	-1.06***	-1.93***	-2.51**
	(0.48)	(0.37)	(0.47)	(1.02)
AR p-value	0.67	0.31	0.06	0.07
Observations	1321	1317	1311	1306

Notes: This Table displays the Phillips Multiplier estimations using the exogenous changes in the short term interest rate as instrument. The sample is annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on Equation 7 via a Panel LP-IV approach. It reports the Anderson and Rubin (1949) p-values for the test that the multiplier estimates differ across states which is robust to weak instruments, coherent to what the F-Statistic displays. The dependent variable was truncated at the top 1% and bottom 1%. Figure A.7 displays the graph counterpart of this Table. Estimates for the Effective F statistic applying the procedure in Olea and Pflueger (2013) for each state independently are in Figure A.8.

Table 5 reports the multipliers in times of low and high price inflation. The state dependent multipliers are statistically different at horizons 2 and 3. In periods of low inflation the long-run trade-off between inflation and unemployment is less exploitable. That is, given this weaker trade-off, central bankers are less able to steer inflation when facing periods of low inflation.

Again, this result does not qualitatively change if we adjust the state definition either by excluding deflation periods from the analysis or by adjusting the upper threshold between 1 and 3%, in fact, the lower the upper bound the more significant the results become.[16]

6 Conclusion

The present paper starts by documenting the changes in the wage Phillips curve using reduced form regressions applied to a panel of advanced economies since 1870. In particular, it provides evidence of a substantial variation in the estimated coefficients on both lagged price inflation and unemployment rate in the wage Phillips curve throughout the sample.

The estimated wage Phillips curves show that most qualitative findings from the reduced form evidence are not driven by endogeneity problems and that the reduced sensitivity of wage inflation to unemployment is also reflected in the estimated changes in a multiplier statistic, based on the estimated impulse responses to monetary policy shocks.

I draw two main conclusions from my findings. First, I confirm the existence of a growing disconnect between wage inflation and unemployment in recent periods, also acknowledging that such disconnect already existed during the Gold Standard. Secondly, I provide evidence in favor of a mechanism that sheds some light on the nature of this phenomenon: a low price inflation environment promotes this disconnect – hence displaying a flatter wage Phillips curve.

[16]Results can be made available from the author.

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Appendix A Supporting figures and tables

Table A.1: Wage Inflation Correlations Table

	π_t^p	π_{t-1}^p	u_t
Australia	0.693	0.688	-0.404
Belgium	0.532	0.677	-0.459
Canada	0.805	0.496	-0.422
Denmark	0.626	0.644	-0.089
Finland	0.378	0.496	-0.429
France	0.863	0.761	-0.515
Germany	0.680	0.629	-0.500
Italy	0.634	0.772	-0.095
Japan	0.310	0.448	-0.749
Netherlands	0.627	0.580	-0.322
Norway	0.793	0.717	-0.600
Portugal	0.686	0.632	-0.232
Spain	0.545	0.465	-0.265
Sweden	0.793	0.754	-0.501
Switzerland	0.603	0.706	-0.438
UK	0.739	0.674	-0.210
USA	0.843	0.548	-0.205

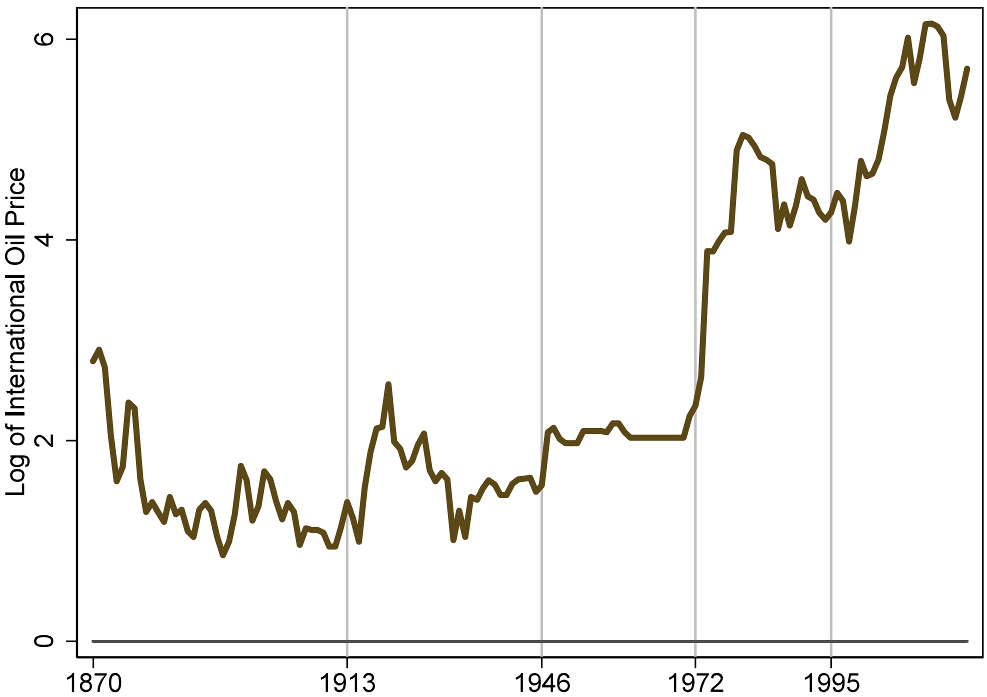
Notes: Correlation between wage inflation and price inflation, lagged price inflation, and unemployment by country in the main sample excluding outliers as defined in the text.

Table A.2: Descriptive statistics - full sample

	N	Mean	Std. Dev.	Min	Max
1870-1913					
Unemployment rate	225	4.08	2.73	0.20	18.40
Wage inflation	730	1.81	4.23	-24.64	23.80
Price inflation	731	0.46	4.76	-26.91	33.31
World Wars					
Unemployment rate	83	3.27	2.50	0.40	12.51
Wage inflation	134	18.50	21.91	-24.98	97.30
Price inflation	134	16.38	28.16	-37.68	241.41
1922-1938					
Unemployment rate	245	7.35	5.09	0.60	24.90
Wage inflation	281	0.89	8.80	-27.72	62.76
Price inflation	284	0.08	7.55	-18.00	57.05
1946-1971					
Unemployment rate	424	2.50	1.78	0.04	9.92
Wage inflation	439	10.21	19.97	-55.42	225.19
Price inflation	442	5.47	10.64	-17.60	125.33
1972-1994					
Unemployment rate	391	6.40	4.00	0.04	24.21
Wage inflation	391	9.23	6.23	-1.08	32.28
Price inflation	391	7.53	5.45	-0.71	37.88
1995-2018					
Unemployment rate	408	7.36	3.73	2.12	26.09
Wage inflation	408	2.51	1.68	-3.40	7.72
Price inflation	408	1.71	1.15	-1.35	5.24
Total					
Unemployment rate	1776	5.38	4.04	0.04	26.09
Wage inflation	2383	5.52	12.07	-55.42	225.19
Price inflation	2390	3.60	10.06	-37.68	241.41

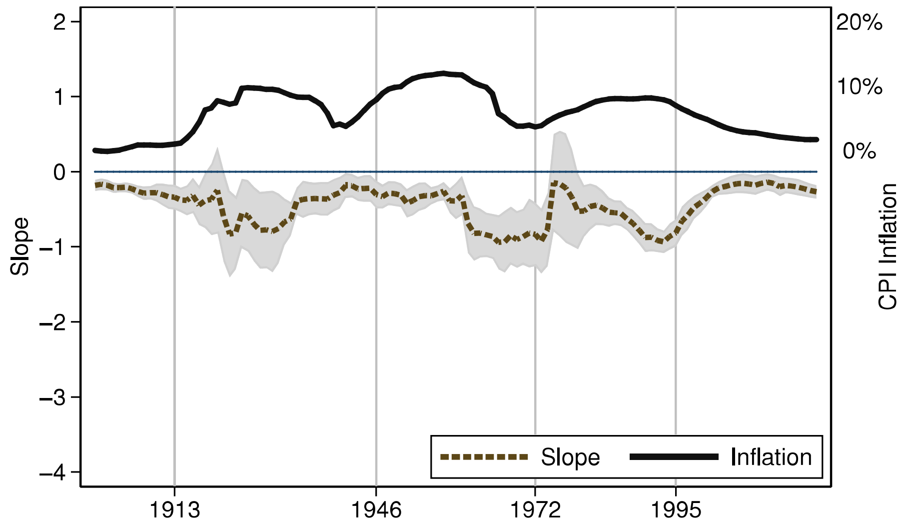
Notes: All values are in percent. The hyperinflation in Germany (1920-1926) is not included. All observations available in the dataset.

Figure A.1: International Oil Price Inflation



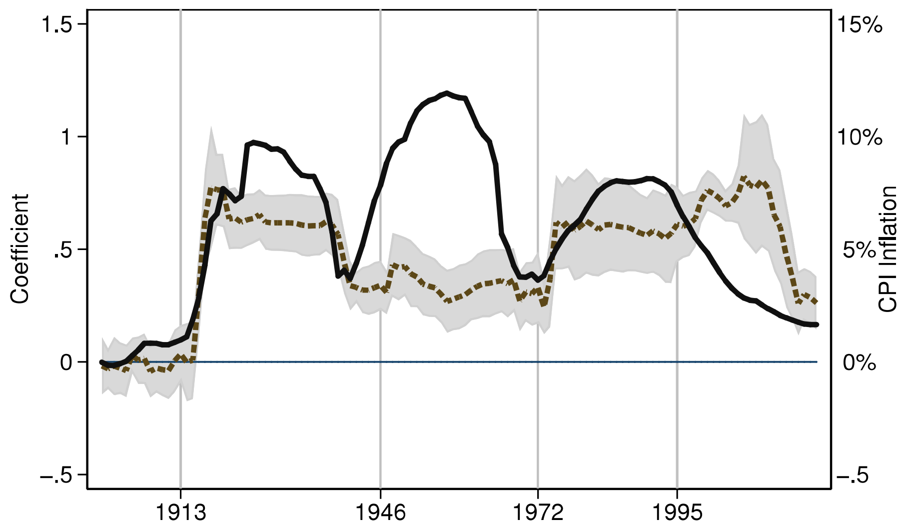
Notes: This figure plots the evolution of the logarithm of the international crude oil price in current US dollars. Sources: from 1870 to 1944 uses a US Average, from 1945 to 1983 uses Arabian Light posted at Ras Tanura, and from 1984 to 2018 uses Brent dated. More details on this report here, consulted on 17th July, 2019.

Figure A.2: OLS 20-year Rolling Window (φ)



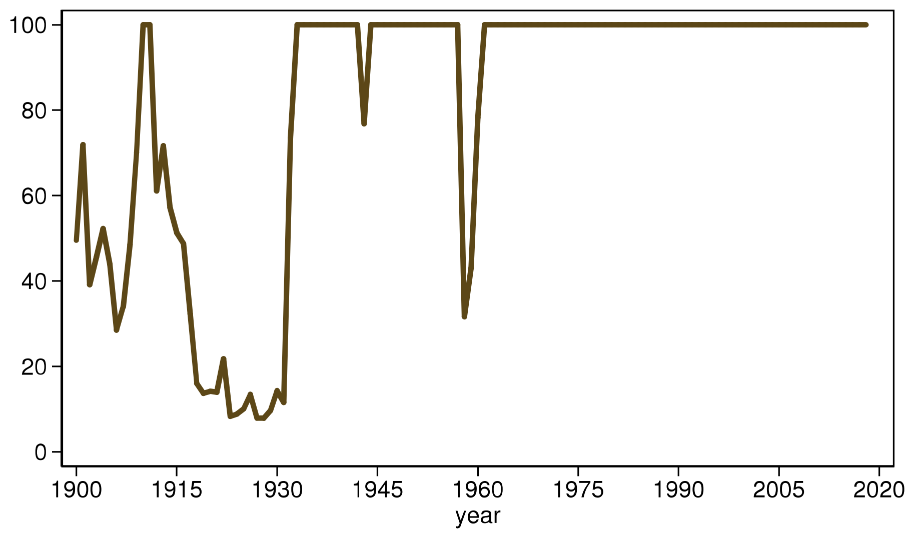
Notes: This figure plots a time-varying estimate of the slope of the wage Phillips curve using OLS and annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on a rolling OLS regression with a 20-year window and a 90% confidence band. The dependent variable was truncated at the top 1% and bottom 1%.

Figure A.3: OLS 20-year Rolling Window (γ)



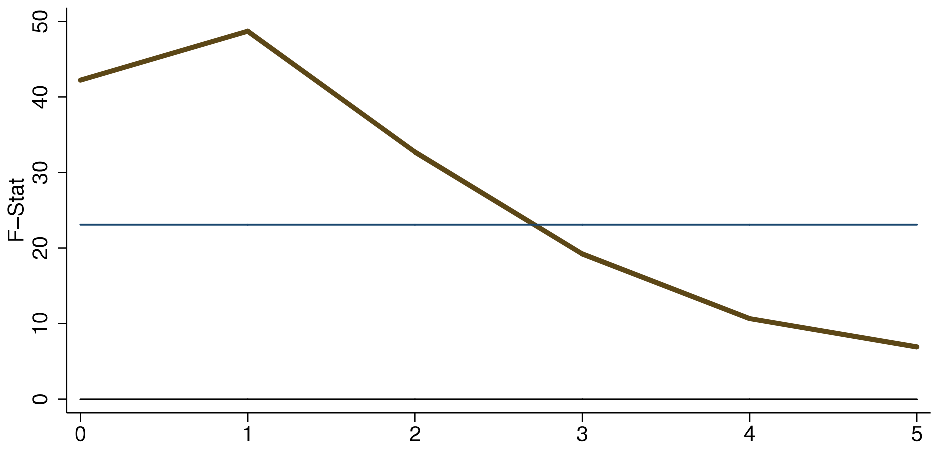
Notes: This figure plots a time-varying estimate of the lagged inflation coefficient (γ) using OLS and annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on a rolling OLS regression with a 20-year window and a 90% confidence band. The dependent variable was truncated at the top 1% and bottom 1%.

Figure A.4: IV 20-year Rolling Window F-Statistic



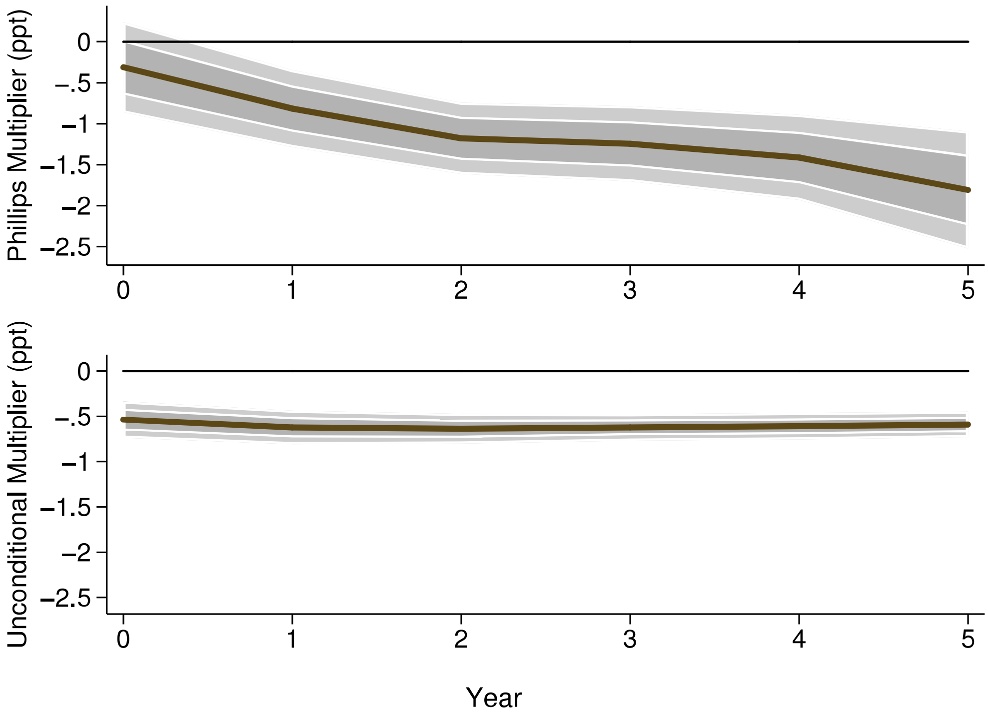
Notes: This figure plots the time-varying estimate of the Kleibergen-Paap Wald rk F-statistic. The estimates were truncated at 100.

Figure A.5: Phillips Multiplier - F-Statistic



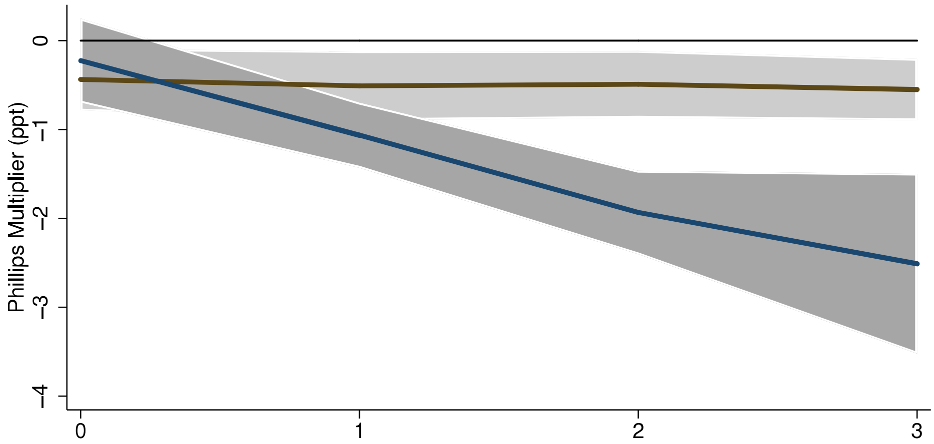
Notes: This figure plots the estimates for the Effective F statistic applying the procedure in Olea and Pflueger (2013). The horizontal line represents the 10% threshold.

Figure A.6: Phillips Multiplier - IV vs OLS



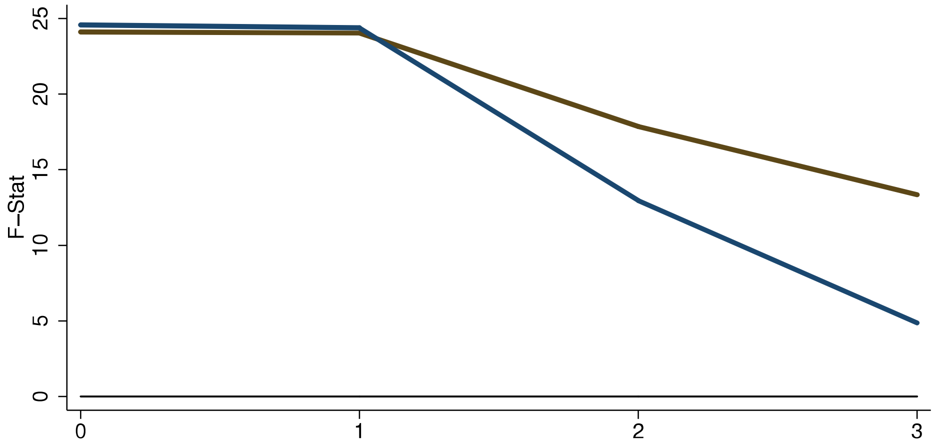
Notes: This figure plots the Phillips Multiplier estimated both by OLS and IV using the exogenous changes in the short term interest rate as instrument. The sample is annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on a Panel LP approach and displays a 68% and 90% confidence bands. The dependent variable was truncated at the top 1% and bottom 1%. Estimates for the Effective F statistic applying the procedure in Olea and Pflueger (2013) are in Figure A.5.

Figure A.7: State Dependent Phillips Multiplier



Notes: This figure plots the Phillips Multiplier for (blue) high and (olive) low inflation periods. They are estimated using the exogenous changes in the short term interest rate as instrument. The sample is annual data from 1870 to 2018 for all 17 countries in the main sample. It is computed based on Equation 7 via a Panel LP-IV approach and displays a 68% confidence band. The dependent variable was truncated at the top 1% and bottom 1%. Estimates for the Effective F statistic applying the procedure in Olea and Pflueger (2013) are in Figure A.8.

Figure A.8: State Dependent Phillips Multiplier - F-Statistic



Notes: This figure plots the estimates for the Effective F statistic applying the procedure in Olea and Pflueger (2013) for (blue) high and (olive) low inflation periods.

Appendix B The New Keynesian wage Phillips curve: a panel framework

In this section, I derive a panel version of the New Keynesian wage Phillips curve following closely the work from Galí (2011) and Erceg, Henderson, and Levin (2000) on staggered wage contracts. Then, I explore the mechanisms which might lead to changes in the Phillips curve slope across time.

As in Galí (2011), the variant of the staggered wage setting model from Erceg et al. (2000) presented here assumes that labor is indivisible. Furthermore, I assume that there are C countries, indexed by c, in order to add a panel component.

In each country, there is a (large) representative household with a continuum of members represented by the unit square and indexed by a pair $(i, j) \in [0, 1] \times [0, 1]$. Where the first dimension represents the type of labor service in which a given household member is specialized and the second one determines the member's disutility from work. The latter is given by:

$$\begin{cases} \chi_{c,t}\ j^{\phi} & if\ employed \\ 0 & if\ unemployed \end{cases}$$

where $\phi \geq 0$ determines the elasticity of the marginal disutility of work and $\chi_{c,t} > 0$ is a country-specific and exogenous labor supply shock which may shift the marginal rate of substitution between consumption and leisure over time.

Utility is logarithmic in consumption and there is full risk sharing among household members. The household period utility is described as:

$$\begin{aligned} U(C_{c,t}, \{N_{c,t}(i)\}, \chi_{c,t}) &\equiv log(C_{c,t}) - \chi_{c,t} \int_0^1 \int_0^{N_{c,t}(i)} j^{\phi} dj di \\ &= log(C_{c,t}) - \chi_{c,t} \int_0^1 \frac{N_{c,t}(i)^{1+\phi}}{1+\phi} di \end{aligned}$$

where $C_{c,t}$ denotes the household consumption, and $N_{c,t}(i) \in [0, 1]$ is the fraction of members specialized in type i labor who are employed in period t in country c. In each country, the representative household maximizes:

$$E_0 \sum_{t=0}^{\infty} \beta^t U(C_{c,t}, \{N_{c,t}(i)\}, \chi_{c,t})$$

subject to a sequence of budget constraints

$$P_{c,t}C_{c,t} + Q_{c,t}B_{c,t} \leq B_{c,t-1} + \int_0^1 W_{c,t}(i)N_{c,t}(i)di + \Pi_t$$

where $P_{c,t}$ is the price of the consumption bundle. $W_{c,t}(i)$ is the nominal wage for labor of type i, $B_{c,t}$ represents purchases of a nominally risk-free one-period bond (at the national price $Q_{c,t}$), and $\Pi_{c,t}$ is a lump-sum component of income.[17]

As in Erceg et al. (2000) and following the formalism of Calvo (1983), workers supplying a labor service of a given type get to reset their (nominal) wage with probability $(1-\theta)$ each period. For now, I assume that this probability is constant over time and across countries just like the two other parameters β and ϕ. In the next section, I delve further into this issue.

Once the wage has been set, the quantity of workers employed is determined unilaterally by firms, with households willingly meeting that demand (given that the wage remains above the disutility of work for the marginal worker). When workers reoptimize their wage in period t, they do so by choosing a wage $W^*_{c,t}$ such that the following first order condition holds:[18]

$$\sum_{k=0}^{\infty}(\beta\theta)^k E_t\left[\frac{N_{c,t+k|t}}{C_{c,t+k}}\left(\frac{W_t^{c*}}{P_{c,t+k}} - \mathcal{M}_c MRS_{c,t+k|t}\right)\right] = 0$$

where $N_{c,t+k|t}$ denotes the quantity demanded in period $t+k$ of a labor type whose wage is being reset in period t, $MRS_{c,t+k|t} \equiv \chi_{c,t+k} C_{c,t+k}(N_{c,t+k|t})^{\phi}$ is the relevant marginal rate of substitution between consumption and employment in period $t+k$, and $\mathcal{M}_c \equiv \dfrac{\epsilon_c}{(\epsilon_c - 1)}$ is the desired (or flexible wage) markup, with ϵ_c denoting the wage elasticity of demand for the services of each labor type assumed to be constant over time.[19]

Log-linearizing the above optimality condition around a perfect foresight zero inflation steady state, using lower case letters to denote the logs of the corresponding variable and letting $\mu_c \equiv log(\mathcal{M}_c)$, we obtain the approximate wage setting rule:

$$w^*_{c,t} = \mu_c + (1-\beta\theta)\sum_{k=0}^{\infty}(\beta\theta)^k E_t[mrs_{c,t+k|t} + p_{c,t+k}] \tag{8}$$

When nominal wage rigidities are present ($\theta > 0$), new wages are set as a constant markup over a weighted average of current and expected future price-adjusted marginal rates of substitution. Furthermore, log-linearizing the expression for the aggregate wage index around a zero inflation steady state we obtain:

$$w_{c,t} = (1-\theta)w^*_{c,t} + \theta w_{c,t-1} \tag{9}$$

[17] As in Galí (2011) the sequence of period budget constraints is supplemented with a solvency condition which prevents the household from engaging in Ponzi schemes.

[18] For details of the optimal wage setting condition derivation see Erceg et al. (2000).

[19] See Galí, Smets, and Wouters (2012) for an application of a one-country framework allowing for time variations in the desired wage markup.

Moreover, letting $mrs_{c,t} \equiv c_{c,t} + \phi n_{c,t} + \xi_{c,t}$ denote the economy's average (log) marginal rate of substitution, with $\xi_{c,t} = log(\chi_{c,t})$, we can write:

$$mrs_{c,t+k|t} = mrs_{c,t+k} - \epsilon_c \phi (w^*_{c,t} - w_{c,t+k}) \tag{10}$$

I proceed to combine equations (1) through (3) in order to derive the baseline wage equation in each country c:

$$\pi^w_{c,t} = \beta E_t[\pi^w_{c,t+1}] + \lambda_c(\mu_{c,t} - \mu_c) \tag{11}$$

where $\pi^w_{c,t} \equiv w_{c,t} - w_{c,t-1}$ is wage inflation, $\mu_{c,t} \equiv w_{c,t} - p_{c,t} - mrs_{c,t}$ denotes the (average) wage markup in country c and $\mu_{c,t} - \mu_c$ is the natural wage markup, the gap between the average real wage and the marginal rate of substitution that would prevail under flexible prices. By assumption, we then have $\lambda_c \equiv -\dfrac{(1-\theta)(1-\beta\theta)}{\theta(1+\epsilon_c\phi)} < 0$.

Even though this relation holds irrespective of the wage setting process, I further assume partial indexation of wages to a measure of lagged price inflation (π^p_{t-1}) when there is no reoptimization of wages, with probability θ each period. Such assumption is reasonable when we consider that there exists a strong positive correlation between wage inflation (π^w_t) and lagged price inflation (π^p_{t-1}) at the country level as showed in Table A.1. Following Galí and Gambetti (2020), this changes equation (11) to:

$$\tilde{\pi}^w_{c,t} = \beta E_t[\tilde{\pi}^w_{c,t+1}] + \lambda_c(\mu_{c,t} - \mu_c)$$

where $\tilde{\pi}^w_{c,t} \equiv \pi^w_{c,t} - [\gamma\pi^p_{t-1} + (1-\gamma)\pi]$ and γ is the degree of indexation. Further assuming that the natural wage markup follows an AR(1) with persistence ρ^w one can re-write the previous equation as:

$$\tilde{\pi}^w_{c,t} = \frac{\lambda_c}{1-\beta\rho^w}(\mu_{c,t} - \mu_c) \tag{12}$$

Next, I introduce unemployment explicitly in the model. Following Galí (2011), an household member (i, j) uses a household welfare criterion to decide whether she finds it optimal to participate in the labor market. She does so whenever the real wage prevailing in her trade is above her disutility from working, that is, if and only if:

$$\frac{W_{c,t}(i)}{P_{c,t}} \geq \chi_{c,t} C_{c,t} j^{\phi}$$

Thus, the marginal supplier of type i labor (employed or unemployed) in country c, which I denote by $L_{c,t}(i)$, is implicitly given by:

$$\frac{W_{c,t}(i)}{P_{c,t}} = \chi_{c,t} C_{c,t} L_{c,t}(i)^{\phi}$$

Log-linearizing and aggregating yields:

$$w_{c,t} - p_{c,t} = c_{c,t} + \phi l_{c,t} + \xi_{c,t} \tag{13}$$

where $l_{c,t} \equiv \int_0^1 l_{c,t}(i)di$ can be interpreted as the model's implied labor force, and $w_{c,t} \equiv \int_0^1 w_{c,t}(i)di$ is the average wage, both expressed in logs. Naturally, the unemployment rate can be defined as:

$$u_{c,t} \equiv l_{c,t} - n_{c,t} \tag{14}$$

Combining equations (13) and (14) with the expression for the average wage markup $\mu_{c,t}$ yields the following linear relation between the wage markup and the unemployment rate:

$$u_{c,t} = \mu_{c,t}\phi^{-1} \tag{15}$$

Let us define the natural rate of unemployment, $u^n_{c,t}$, as the rate of unemployment that would prevail in the absence of nominal wage rigidities. It follows from the assumption of a constant desired wage markup that $u^n_{c,t}$ is constant and given by:

$$u^n_c = \mu_c \phi^{-1} \tag{16}$$

Thus, in the framework above unemployment is a consequence of workers' market power (that is, of the wage being above their perfectly competitive level), while unemployment fluctuations result from the slow adjustment of wages.

Finally, combining equations (12), (15) and (16) and defining $\varphi_c \equiv \dfrac{\phi\lambda_c}{1-\beta\rho^w}$ we obtain the following panel version of a New Keynesian wage Phillips curve (henceforth, NKWPC):

$$\pi^w_{c,t} = (1-\gamma)\pi + \gamma\pi^p_{t-1} + \varphi_c(u_{c,t} - u^n_c) \tag{17}$$

B.1 A time-varying New Keynesian wage Phillips curve

A classical assumption across existing studies on the Phillips curve is to have time-invariant parameters $(\beta, \theta, \epsilon, \phi)$. Notwithstanding, one should be wary of it when estimating the Phillips curve using such long samples as this current work. An economy arguably changes within 150 years.

Romer (1990) argued that the frequency of price (wage) adjustments was likely to vary in response to the volatility of the economic environment. In particular, he claimed that the frequency would be lower in an environment of low inflation with lower volatility. This belief is corroborated by the recent empirical works of E. Gagnon (2009) and Nakamura, Steinsson, Sun, and Villar (2018).

Some literature followed this assumption either by modelling time-varying parameters as random variables (Romer (1990), Primiceri (2005)) or by allowing for different parameters in different subsample periods (Galí and Gertler (1999), Smets and Wouters (2007)).

In this work, given its historical perspective, I assumed that the model parameters might vary across subsamples and make use of rolling windows estimations. From Equation 17, one can see how the slope responds to the parameters by looking at its definition:

$$\varphi \equiv -\frac{\phi(1-\theta)(1-\beta\theta)}{(1-\beta\rho^w)\theta(1+\epsilon\phi)} < 0$$

The slope of the NKWPC is negative by assumption. The relation between the slope and the remaining parameters can be found in Figure B.1. Given the support of its fundamental parameters, it depends positively and strongly on θ.

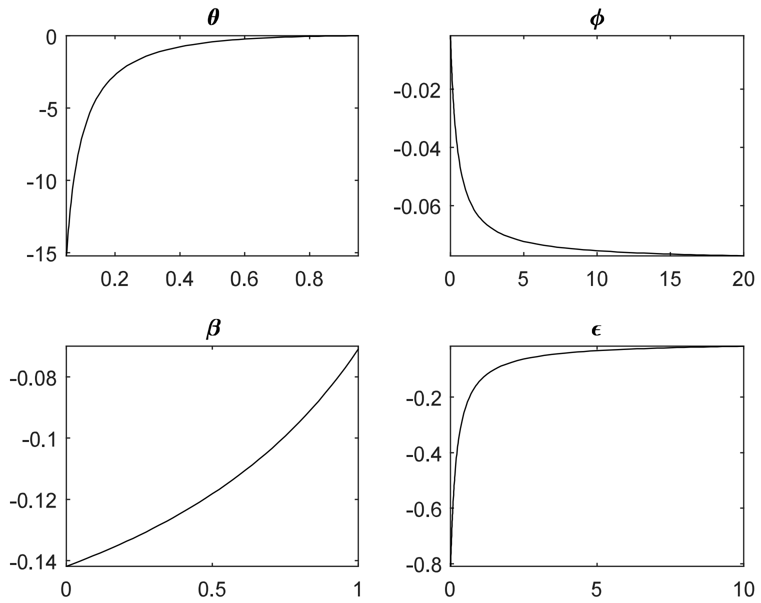
If firms adjust prices more frequently, the NKWPC slope decreases in absolute value. Therefore, one expects that in periods of lower inflation, firms adjust prices and wages less often thus, flattening the slope of the Phillips curve (Benati (2007)). This rationale is in line with two strands of the literature.[20]

The theoretical state dependent pricing literature states that, at low inflation levels, the frequency of wage (price) adjustment decreases. Since workers and firms respond less to shocks at low inflation, the overall price level also becomes less reactive, leading to a substantially flatter Phillips curve (Costain et al. (2019)).

An alternative argument hinges on the theory of nominal price rigidities (Tobin (1972), Akerlof et al. (1996), Benigno and Ricci (2011), Daly and Hobijn (2014), J. Gagnon and Collins (2019), Lindé and Trabandt (2019)). It claims that during periods of low inflation, nominal wages (prices) tend to be more rigid downwards than upwards and thus, will react

[20] It is not the focus of this paper to test which theory is more suitable but rather test whether in periods of low inflation we see a flatter Phillips curve.

Figure B.1: Slope of the NKWPC



Notes: Calibration. The discount factor, β is set to 0.99 and the persistence of the natural wage markup gap ρ^w is set to 0.5. The Calvo wage stickiness parameter, θ, imply an average duration of individual wages of one year, in a way consistent with much of the micro evidence Nakamura and Steinsson (2008). The curvature of labor disutility, ϕ, is set to be 5, a value consistent with a Frisch labor elasticity of 0.2. Galí (2011) shows that ϕ, ϵ, and the steady-state unemployment rate u have the following relationship: $\phi u = log(\epsilon/(\epsilon - 1))$. Given that ϕ is set to 5, the value of ϵ is set to 2.15. This is consistent with a steady-state unemployment rate of 5.4 percent, the average unemployment rate in the full sample (Table 2).

less to changes in the unemployment rate.

Appendix C Historical periods

It is important to have clearly and coherently identified historical periods as we proceed to a deeper investigation of the research question. Table C.1 underlies this analysis. It summarizes the allocation of countries to historical periods and its sources. Given that the main channel being analyzed is the low inflation environment, it is crucial to know when countries were targeting either the price of gold or some consumer price index.

Identifying the periods of the *Gold Standard*, the *Bretton Woods* and the *Explicit Inflation Targeting* from the others was straightforward. It was based on documented evidence on which countries were participating in one of the above mentioned regimes by either keeping their exchange rate fixed to the price of gold or by targeting inflation.

Dates for the Gold Standard come from Reinhart and Rogoff (2009) and can be confirmed using different sources such as Bordo and Rockoff (1996) and Diebold, Husted, and Rush (1991). Dates for the Bretton Woods and the Explicit Inflation Targeting come from Central Banks' websites and were complemented by data from Ilzetzki, Reinhart, and Rogoff (2019). Even though the countries present in the Bretton Woods system agreement (1944) started to progressively adopt a fixed exchange rate, I define the start of the Bretton Woods era in 1946 after the creation of the International Monetary Fund (IMF) in December, 1945.[21] The Bretton Woods system broke down on 15 August, 1971 thus, I decided to classify 1971 as the last effective year of this epoch.

The first World War went from on July 1914 to November 1918 while the second World War started on September, 1939 and ended on September, 1945. Consequently, it was easy to define the *Interwar Period* with the caveat of choosing 1922 as the beginning of this period to remove the effects of the post-war recession mostly felt in Europe, a common practice in the literature.

Finally, the hardest part was to argue on the exact year when countries started implicitly targeting inflation. While explicit inflation targeting, in the sense of it being announced by the national Central Banks, started only in the nineties, there is a long literature arguing that some countries were implicitly doing it before that.

[21] The starting dates for most countries coincide with the effective date of IMF membership available online here. Nevertheless, I acknowledge that if a country had its currency pegged to a major currency such as the US dollar, it is expected that they were implicitly "part of" the Bretton Woods system.

Table C.1: Allocation of countries to historical periods

	Gold Standard	Bretton Woods	Implicit Inflation Target	Explicit Inflation Target	CB Foundation
Australia	1870-1915 Reinhart and Rogoff (2009)	1949-1971 Reserve Bank of Australia	1993-2020 Reserve Bank of Australia	1993-2018 Reserve Bank of Australia	1911
Belgium	1878-1914 RR (2009)	1946-1971 National Bank of Belgium	1990-2020 National Bank of Belgium	no	1850
Canada	1870-1914 RR (2009)	1946-1950 / 1962-1971 Bordo et al. (2010)	1991-2020 Bank of Canada	1991-2020 Bank of Canada	1934
Denmark	1876-1917 RR (2009)	1946-1971 Danmarks Nationalbank	1986-2018 Danmarks Nationalbank	no	1818
Finland	1877-1914 RR (2009)	1948-1971 Suomen Pankki	1995-2020 Suomen Pankki	no	1811
France	1878-1914 RR (2009)	1946-1971 Banque de France	1986-2020 Banque de France	no	1800
Germany	1871-1914 RR (2009)	1952-1971 Bundesbank	1986-2020 Von Hagen (1999)	no	1876
Italy	1884-1917 RR (2009)	1947-1971 Banca D'Italia	1986-2020 Banca D'Italia	no	1893
Japan	1897-1917 RR (2009)	1952-1971 Shizume (2018)	1987-2020 Jinushi et al. (2000)	2012-2020 Bank of Japan	1882
Netherlands	1875-1914 RR (2009)	1946-1971 De Nederlandsche Bank	1986-2020 De Nederlandsche Bank	no	1814
Norway	1875-1914 RR (2009)	1946-1971 Norges Bank	1990-2020 Norges Bank	2001-2020 Norges Bank	1816
Portugal	1870-1891 RR (2009)	1961-1971 Bordo and dos Santos (1995)	1992-2020 Banco de Portugal	no	1846
Spain	— RR (2009)	1958-1971 Banco de España	1990-2020 Banco de España	no	1874
Sweden	1873-1914 RR (2009)	1951-1971 Sveriges Riksbank	1991-2020 Sveriges Riksbank	1993-2020 Sveriges Riksbank	1668
Switzerland	1878-1914 RR (2009)	1946-1971 Baltensperger and Kugler (2017)	1986-2020 Swiss National Bank	no	1907
UK	1870-1914 RR (2009)	1946-1971 Bordo and Schwartz (1999)	1991-2020 Bank of England	1992-2020 Bank of England	1694
USA	1880-1917 RR (2009)	1946-1971 Bordo and Schwartz (1999)	1988-2020 Goodfriend (2004)	2012-2020 FED	1913

Notes: This Table presents the allocation of countries to historical periods. Further discussion on this follows. All Central Bank Foundations dates were collected from the Central Banks' websites.

According to Von Hagen (1999), "*the Bundesbank began to announce inflation targets together with adopting monetary targeting, first a series of 'unavoidable' inflation rates and, from 1986 on, a fixed rate of 2%*".[22] The German Mark accounted for roughly one third of the weight to the European Currency Unit (ECU) value from 1984 until 1999. Given both Germany's importance to ECU and its adoption of an inflation target I am considering that countries who already belong to the European Exchange Rate Mechanism (EERM) such as Denmark, France, Italy, Netherlands and then countries who later join the EERM as Belgium, Finland, Portugal, Spain and the UK or pegged their currencies to the ECU such as Norway and Sweden, are implicitly adopting a behaviour of targeting price inflation. For Japan, this definition is based on the work of Jinushi et al. (2000) who thoroughly explain why Japan was implicitly targeting inflation by 1987.

Ungerer (1997), page 183, tells us that Norway and Sweden pegged its currency to the ECU and in 1990 and 1991 respectively. Even though they abandoned such peg later on, both of them became explicit inflation targeters - Norway in 2001 and Sweden in 1993.

UK was a member of the ERM from October 1990 to September 1992. Shortly after, it became the first explicit inflation targeter in Europe. Finally, and even though only in 2012 did the FED announce its explicit target to inflation, Goodfriend (2004) argues that *"monetary policy as conducted in the Greenspan era can be characterized as implicit inflation targeting"*, hence, I consider that in 1988 the USA was an implicit targeter.

According to Berg and Jonung (1999), Sweden had an experience similar to inflation targeting between 1931-1937 thus undergoing into an attempt to mimic the price stabilization features of the gold standard while eliminating the volatility produced by shocks to the gold market.

[22] Some others would agree on a date even before 1986 (Mishkin, 2001).