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The Neutrality of Nominal Rates: How Long is the Long Run?

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Abstract
How can inflation be raised in economies such as Japan and the euro area where it has been below the objective for quite some time? We estimate an empirical model aimed at identifying the effects of permanent and temporary monetary shocks for the U.S., Japan, France, the U.K., Germany and the euro area. We find that the permanent monetary shock leads to a permanent rise in nominal rates and inflation. Importantly, the short-run effects of this permanent shock are similar to the long-run effects: inflation responds positively and immediately to a permanent rise in nominal rates, confirming the results in Uribe (2017, 2018). We also reinvestigate the long-run relation between inflation and nominal short interest rates. Using data for 41 developed countries covering the last 50 years, we document a strong, yet below one-for-one relationship between nominal rates and inflation, that tends to be less visible over the more recent period, characterized by inflation targeting at low common levels.

JEL: E31, E32, E52, E58

Keywords: Fisher Relation, Monetary Neutrality, Panel Cointegration, Monetary Shocks, Structural VECM Models.

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1. Introduction

Inflation has been low in Japan for 25 years. Average CPI inflation since the 1990s has been 0.5%, clearly below the objective that in 2013 was raised to 2%.[1] The low inflation has not been associated with low economic growth or high unemployment that was 2.4% in April of 2019. Policy in Japan has tried every possible stimulus, both fiscal and monetary. From the early 1990s to the global financial crisis period about 15 fiscal packages were introduced. Fiscal deficits have been high and rising and the resulting gross public debt was at the end of 2018 close to 250% of GDP (60% in the early 1990s).[2] The Bank of Japan (BoJ) cut policy rates to zero in the 1990s, and has kept them there for most of the last 25 years. Since the early 2000s, the BoJ has also used numerous other monetary policy measures aimed at raising inflation, including quantitative easing, the higher target for inflation, policies aimed at lowering long-term rates, an inflation-overshooting commitment and forward guidance. The size of the BoJ balance sheet has increased to close to one GDP. None of these measures have been able to bring inflation close to target.

The euro area is starting to look like Japan. Since 2013, inflation in the euro area has been well below 2% most of the time. Average HICP inflation in the last 10 years is 1.3% (1.1% core). The policy rate has been close to zero since 2014 and is expected to remain there. There has also been ample experimentation with unconventional balance sheet policies. It is hard to argue that current low inflation in the euro area is related to economic slack. The unemployment rate in April of 2019 was 7.7%, close to the historical minimum, down from 12% in early 2014.

How can inflation be raised in Japan and possibly also in the euro area? This paper is an empirical assessment of a policy that aims at raising inflation by raising policy rates. The mechanism is orthogonal to that behind the standard policy of stimulating the economy by reducing rates. It hinges on the long-run neutrality of money. In short, if nominal rates stay persistently low, it cannot be that real rates are persistently kept low, but rather that inflation cannot be high. This is an old argument, notably in Fisher's "*Appreciation and Interest*" (1896), but dating back at least to the 1740s (see Humphrey (1983) for an historical review).[3] More recently, given the prolonged experience with low nominal interest rates, a revived strand in the literature has argued that such policy can lead to lower inflation, see Cochrane (2016), Garín, Lester and Sims (2018) or the analysis in Benhabib, Schmitt-Grohé and Uribe (2001). These papers are references for the so-called

1. Inflation went up on two occasions with increases in the VAT rate that were reversed after one year.

2. The net debt was 150% of GDP.

3. Fisher realizes that over the short run changes in nominal rates or inflation could affect real rates, and suggests mechanisms for the adjustment towards the "equilibrium" real rate, while his description of the factors affecting this "natural" rate certainly makes plausible that it changes over time. See also Fisher (1930).

"Neo-Fisherian" explanations of the low inflation outcomes experienced in Japan, Europe, and even the U.S. during the zero lower bound episodes.

The first part of the paper is a re-investigation of the relation between inflation and nominal short rates.[4] We start by considering standard New Keynesian models, where the Fisher relation is postulated, in order to discuss the empirical challenges to that task. We analyze regression coefficients in models featuring either a varying inflation target, displaying a unit-root and assumed to be cointegrated with nominal rates, or a fixed inflation target. Regardless of the importance of the various shocks in the non-stationary model, data generated by it would allow the estimation of a one-for-one relation between inflation and nominal rates, even if the precision can be low if the shocks to the target are small. Turning to the stationary (fixed target) model, the regression coefficient of inflation on nominal rates can be pretty much anything.[5]

We proceed by considering a panel of countries with data from 1960 to 2016. We provide several insights. First, we show that a cross-section of averaged data for the full sample is unambiguous about a positive, close to one-for-one relationship between the two variables. Since the later part of the sample is a period of successful inflation targeting, we break the sample in two and find that a simple estimation of this relationship yields a lower coefficient in that part of the sample. Still, the fit to the line with unit slope is as good as the one obtained using the whole sample. This could mean that the one-for-one relation is a feature of the data, but harder to detect if (many) countries successfully control inflation around a common level; the estimated slope could well be zero.

We then turn to a time series and panel cointegration analysis, which reveals similar evidence. For the whole sample, characterized by non-stationarity in inflation and nominal rates we do estimate a slope close to, but below, unity. For the later part of the sample, one better characterized by stationary inflation and nominal rates, the slope coefficients are substantially lower than one. Still, it should be noted that the residual dispersion around a unit slope line is lower than that obtained over the whole sample.

The central part of the paper follows. We estimate the effects of a permanent monetary policy shock, employing a very standard Vector Error Correction model for the U.S., Japan, France, the U.K., Germany and the euro area, using monthly data up until 2018. We argue that there is strong evidence for a positive permanent monetary shock, corresponding to higher nominal rates, to raise inflation not only in

4. Investigating the Fisher relation has obviously been done before, see e.g. Mishkin (1992), Crowder and Hoffman (1996), Koustas and Serletis (1999) or Westerlund (2008) in a panel setting for more recent references.

5. These difficulties in testing long-run relations, or neutrality, using reduced form regressions are obviously well-known since at least Lucas (1972) and Sargent (1973). Knowledge of the complete economic model can often be dispensed with if the data is characterized by permanent shifts, likewise a unit-root behavior in the relevant time series, see Fisher and Seater (1993), King and Watson (1997) and other references therein. Results in this paper ought to be interpreted with these tensions in mind.

the long run but also in the short run. Our empirical model allows us to characterize how long it takes for inflation and nominal rates to adjust towards the long-run relation while allowing for the presence of temporary monetary shocks. Identification of the two nominal shocks exploits the cointegration relation and requires some further (mild) identifying assumptions, namely that permanent shocks to output do not affect the level of inflation and nominal rates in the long run; this is perhaps the simplest and most innocuous way of empirically identifying the two types of nominal shocks.

We find that the effects of temporary nominal shocks are the standard ones, i.e., a temporary rise in nominal rates leads to a temporary fall in inflation; the effects dissipate in 3 years. More interesting is the effect of permanent nominal shocks. Given cointegration between inflation and nominal rates - whose coefficient we do not impose to be one -, permanent shocks lead to permanent shifts in nominal rates and inflation; after three years the adjustment is essentially made. Again, the short-run effects are similar to the long-run effects in that inflation responds positively at short horizons to a permanent rise in nominal rates. Qualitatively, the impulse responses we obtain in our empirical model are remarkably similar to what obtains in the version of the New Keynesian model with a unit-root in inflation and nominal rates.

We confirm the results in Uribe (2017, 2018), who employs a different modeling strategy, applied to the U.S. and Japan. He imposes a unitary cointegration coefficient and assumes, for identification, the standard effects of temporary shocks on impact.[6] Further, as in our model, permanent shocks are found to account for the bulk of the forecast error variance decomposition in inflation and nominal rates at horizons above (and often below) 2 years. It seems thus critical to distinguish between permanent and temporary monetary shocks in interpreting the data.

All in all we conclude that the Fisher relation is firmly established in the data and reveals itself long before the long run.

The outline of the paper is as follows. Section 2 sketches the theoretical one-for-one long-run relation between nominal rates and inflation and highlights the difficulties in identifying empirically this relation, analyzing regression coefficients obtained in New-Keynesian models featuring either a fixed target for inflation or a unit-root in the target along with cointegration between inflation and nominal rates. Section 3 looks at the cross-sectional, time series and panel evidence. Section 4 discusses results from an identified structural Vector Error Correction model. Section 5 concludes. The data and details of the models used can be found in the Appendix. An extended Supplemental Material Appendix provides further results.

6. Fève, Matheron, and Sahuc (2010) take an approach similar to ours. They focus on disinflation policy in Europe before the euro, estimate a structural VAR and identify a permanent nominal shock as the only shock affecting nominal rates over the long-run. Cointegration between inflation and nominal rates with unitary coefficient is also imposed. Due to the focus on the long run relation no attempt is made to distinguish between permanent and temporary nominal shocks.

2. The Fisher Relationship and its Estimation

The Fisher relation is established using a simple no-arbitrage argument: loans in terms of goods or in terms of money should cost the same, i.e. $(1+R) = (1+\pi)(1+r)$, where R is a nominal rate, π is expected inflation and r is the real rate on loans made in terms of goods. The important claim is that r is fixed and independent of R or π, a "natural" rate. The implication of this neutrality proposition is that R and π move one-for-one. Fisher's description of the factors affecting this "natural" r certainly makes plausible that it changes over time, see also Fisher (1930).

This relationship is embedded (postulated) in the core of modern macroeconomic models. Under certain conditions, along a steady-state growth path the following holds after linearization:

$$R_t - \pi_{t+1} = 1/\beta - 1 + \gamma \tag{1}$$

where we keep the subscript t for generality. The right-hand side is the (exogenous) real rate, where β is a discount factor and γ is the growth rate of total factor productivity. That is, the nominal interest rate minus inflation equals a real rate that is exogenous to monetary policy. Hence, on average, for a given steady-state real interest rate, a higher nominal interest rate will necessarily be accompanied by a higher inflation rate. Depending on other specificities, such as taxes, the expression for this natural real rate may have some other elements, but the fact that it is exogenous to monetary policy does not change in different models. Away from the steady state, changes in inflation or nominal rates affect real rates in several formulations. Now, notice that it is the steady state of $R_t - \pi_{t+1}$ that is constant. If there is a steady state for π_{t+1} (say, a fixed inflation target) then there is a steady state for R_t. But inflation does not need to be stationary around this steady state. In that case R_t will not be stationary but $R_t - \pi_{t+1}$ will be stationary. In time series jargon, inflation and nominal rates would be cointegrated, if the form of non-stationarity is a unit-root in the two series. It is straightforward to consider such complication in the theoretical models, see Section 2.1 for one possibility.[7]

This characterization turns out to be very relevant if one is interested in confirming empirically that a Fisher relation characterizes the data. In fact, if inflation and nominal rates are stationary, failure to find a unitary coefficient using (contemporaneous or future) inflation and interest rate data using regressions or other reduced form methods says little about the validity of the Fisher relation. That identification would require a complete knowledge of the model generating

7. Needless to say, complications would arise if the natural real rate is not stationary, e.g. due to changes in demographics or in the growth rate of technology (some sort of secular stagnation), but these changes are bound to be slow and a stationary real rate might be a good approximation. In this case the difference $R_t - \pi_{t+1}$ would not be stationary in the data and a more realistic model would reflect this feature. That is seldom done in the literature, which typically postulates that $R_t - \pi_{t+1}$ has a steady state, and most often that R_t **and** π_{t+1} have a steady state.

the data. This is just a restatement of the general points made in Lucas (1972) or Sargent (1973). In order to see this, take the standard three equations sticky prices model - a Fisher equation, a Phillips curve and a Taylor rule -, log-linearized around the steady-state:

$$y_t = E_t(y_{t+1}) - \sigma(R_t - E_t(\pi_{t+1})) + \varepsilon_t^{\beta} \tag{2}$$

$$\pi_t = \beta E_t(\pi_{t+1}) + \lambda y_t + \varepsilon_t^{c} \tag{3}$$

$$R_t = \varphi_\pi \pi_t + \varepsilon_t^{m} \tag{4}$$

where y_t is output and the $\varepsilon' s$ denote either a preference shock (ε_t^{β}), a cost-push shock (ε_t^{c}) or a monetary shock (ε_t^{m}), assumed to be serially and mutually uncorrelated.[8] In this case it can be shown that a regression of inflation on nominal rates yields the following slope:

$$\frac{Cov(R_t,\pi_t)}{Var(R_t)} = \frac{-\lambda\sigma.Var(\varepsilon_t^{m}) + \varphi_\pi.Var(\varepsilon_t^{c}) + \lambda^2\varphi_\pi.Var(\varepsilon_t^{\beta})}{Var(\varepsilon_t^{m}) + \varphi_\pi{}^2.Var(\varepsilon_t^{c}) + \lambda^2\varphi_\pi{}^2.Var(\varepsilon_t^{\beta})} \tag{5}$$

which is necessarily less than 1. If there are only monetary shocks the slope will be negative, equal to $-\lambda\sigma$. If there are only cost push shocks or only preference shocks the slope will be positive, equal to $1/\varphi_\pi$. Price frictions and diverse sources of shocks can thus result in a wide range for the regression coefficient, even a negative one if monetary shocks are the main source of variation in the data.[9] Table 1 reports some values of this slope for a standard parameterization, varying the number of shocks and their relative importance. The slope is bounded above by $1/\varphi_\pi$ which in this case is 2/3. As φ_π approaches one (from above) the slope would approach one. Again, the higher the variance of the monetary shock the lower is the coefficient; it can even be somewhat negative.

Failure to estimate a unitary coefficient is not due to sticky prices. Take the simplest model with monetary neutrality - a Fisher relation and a Taylor rule, see Cochrane (2011):

$$R_t = r + E_t(\pi_{t+1}) \tag{6}$$

$$R_t = r + \varphi_\pi \pi_t + \varepsilon_t \tag{7}$$

8. σ is the inverse of the intertemporal elasticity of substitution, β is agents' discount rate, φ_π is the coefficient of inflation in the Taylor rule (assumed greater than 1) and λ is the coefficient of output in the Phillips curve, which is a function of the degree of price stickiness and other deep parameters.

9. The characterization of this slope coefficient is more cumbersome if there is persistence in the shocks, but the main message is unaltered: obtaining a unitary coefficient occurs only in special cases. Also, even without persistence in the shocks, a regression of π_{t+1} on R_t yields a zero coefficient, since π_{t+1} depends only on shocks at $t+1$, which are independent of time t variables. Persistence in the shocks would yield a non-zero coefficient, but again, in general different from one.

where r is the exogenous real rate. The closed-form solution to inflation in this model is given by:

$$\pi_t = -\sum_{j=0}^{\infty} \frac{1}{\varphi_\pi{}^{j+1}} E_t(\varepsilon_{t+j}) \tag{8}$$

Assuming that $\varepsilon_t = \rho\varepsilon_{t-1} + \varepsilon_t^m$, where $\{\varepsilon_t^m\}$ is i.i.d. with mean zero, $|\rho| < 1$ and $\varphi_\pi > 1$ it follows that:

$$E_t(\pi_{t+1}) = \rho\pi_t \tag{9}$$

Therefore, $R_t = r + \rho\pi_t$. A regression of inflation on nominal rates will yield $1/\rho$, the inverse of the persistence of the monetary shock ε_t, and a perfect fit.[10]

What is required in general to identify the Fisher relation using reduced form methods (i.e. without a detailed knowledge of the structural model generating the data) are permanent shifts in the two variables, as stressed by King and Watson (1997) more generally. Below we investigate the slope of a regression of inflation on nominal rates upon introducing, within the standard New Keynesian model, a unit-root in the two variables while keeping the real interest rate stationary (i.e., imposing a cointegration relation between the two variables).

2.1. Introducing a varying inflation target

We introduce a unit-root in inflation and nominal rates in the standard New-Keynesian model by specifying a varying inflation target. We follow closely Juillard et al. (2008), see also Ireland (2009). Details on the microfoundations and extra auxiliary equations can be found in the Appendix. A pricing decision mechanism slightly different from Calvo pricing is employed. Whereas in each period each firm faces a probability of not being able to change its last decision and reoptimize (a la Calvo), the decision is not only about which price to charge but also the growth rate of its price change (every period) until it is able to reoptimize again. The model is therefore less *ad hoc* than one imposing a price indexation mechanism. Importantly, the inflation target, $\overline{\pi}_t$, is now assumed to evolve according to $\overline{\pi}_t = \overline{\pi}_{t-1} + \varepsilon_t^{\overline{\pi}}$ where $\varepsilon_t^{\overline{\pi}}$ is a white-noise shock to the inflation target, assumed uncorrelated with the other shocks in the model. The model retains a Fisher equation and a slightly more complicated Phillips curve. Real rates are still assumed to be stationary; in order to solve the model the gross nominal interest rate $(1 + R_t)$ and gross inflation (Π_t) are thus stationarized as $(1 + R_t)/\overline{\Pi}_t$ and $\Pi_t/\overline{\Pi}_t$, which justifies an additional term

10. In this particular case a regression of π_{t+1} on R_t would find the unitary slope as the expectation error ($\upsilon_t := \pi_{t+1} - E_t(\pi_{t+1})$) is on average 0 and independent of R_t. Therefore $\pi_{t+1} = E_t(\pi_{t+1}) + \upsilon_t = \rho\pi_t + \upsilon_t = R_t - r + \upsilon_t$, allowing identification of the unitary coefficient on R_t. But if, say, $R_t = r + E_t(\pi_{t+1}) + \varepsilon_t^\beta$ and inflation is exogenous and driven by $\pi_t = \rho\pi_{t-1} + \varepsilon_t^c$, with $|\rho| < 1$, ε_t^β and ε_t^c independent and i.i.d. with mean zero (see, e.g., McCallum 1984) then the regression of π_t on R_t yields some $\mu < 1$ (unless the variance of the shock ε_t^β is nil), and a regression of π_{t+1} on R_t yields $\rho\mu$.

1 Shock	σ^{β}	σ^{c}	σ^{m}	Coeff.
	1	0	0	0.66
	0	1	0	0.66
	0	0	1	-0.47
2 Shocks				
	0.5	0.5	0	0.66
	0.5	0	0.5	-0.09
	0	0.5	0.5	0.32
3 Shocks				
	0.4	0.3	0.3	0.39
	0.3	0.4	0.3	0.46
	0.3	0.3	0.4	0.22
	0.7	0.2	0.1	0.63
	0.2	0.7	0.1	0.66
	0.1	0.2	0.7	-0.29

Parameter setting: $\sigma = 1; \beta = 0.99; \lambda = 0.47; \varphi_\pi = 1.5$. σ^{β}, σ^{c} and σ^{m} are the standard deviations of the shocks.

Table 1. Regression Coefficient of π_t on R_t in 3-equation NK model

$\overline{\Pi}_t$ in the Taylor rule:

$$1 + R_t = (1 + R_{t-1})^{\rho}[(\frac{\Pi_t}{\overline{\Pi}_t})^{\varphi_\pi} \frac{1+\gamma}{\beta} \overline{\Pi}_t (\frac{Y_t}{Y_t^*})^{r_y} (\frac{Y_t}{Y_t^*} / \frac{Y_{t-1}}{Y_{t-1}^*})^{r_{\triangle y}}]^{1-\rho} \exp(\varepsilon_t^m)$$

where Y_t is output, Y_t^* is potential output (i.e., output under flexible prices) and ε_t^m is a monetary policy shock. β is the discount factor of households, γ is the growth rate of the economy in the steady-state and ρ, r_π, r_y and $r_{\triangle y}$ govern the reactions of the monetary authority. In log-linearized form this reads as $\widehat{R}_t = \rho(\widehat{R}_{t-1} - \varepsilon_t^{\overline{\pi}}) + (1-\rho)(r_\pi \widehat{\pi}_t + r_y(y_t - y_t^*)) + r_{\triangle y}((y_t - y_t^*) - (y_{t-1} - y_{t-1}^*)) + \varepsilon_t^m$ where $\widehat{R}_t$, $\widehat{\pi}_t$ are log deviations of $(1+R_t)/\overline{\Pi}_t$ and $\Pi_t/\overline{\Pi}_t$ from their respective steady-states and the remaining lower cases denote log deviations from the steady-state. The permanent shock shows up directly in the Taylor rule; it can thus be interpreted as a permanent shift in the average level of nominal rates which comes with a new (implicit) inflation target.

We want to run regressions of inflation $(\Pi_t - 1)$ on nominal interest rates (R_t) using data generated by this model.[11] For our illustrations we use the common parameters of the three equations model of the previous section, restrict ρ and

11. It is straightforward to map nominal interest rates and inflation (which are not variables in the solution to the model) to variables in the model. Defining $1 + R_t^+ := (1+R_t)/\overline{\Pi}_t$, notice that $\log(1+R_t) - \log(1+R_{t-1}) = \log(\overline{\Pi}_t(1+R_t^+)) - \log(\overline{\Pi}_{t-1}(1+R_{t-1}^+)) = \log(\overline{\Pi}_t) - \log(\overline{\Pi}_{t-1}) + \log(1+R_t^+) - \log(1+R_{t-1}^+) = \varepsilon_t^{\overline{\pi}} + \widehat{R}_t - \widehat{R}_{t-1}$. Similarly, $\log(\Pi_t) - \log(\Pi_{t-1}) = \widehat{\pi}_t - \widehat{\pi}_{t-1} + \varepsilon_t^{\overline{\pi}}$ allowing us to simulate realizations of R_t and Π_t.

r_y as well as $r_{\triangle y}$ to zero in the modified Taylor rule and employ the standard deviations of the shocks ε_t^{β}, ε_t^{c} and ε_t^{m} estimated in Valle e Azevedo and Jalles (2017) using U.S. data at a quarterly frequency. While it is clear that the postulated cointegration relation delivers a regression coefficient (of inflation on nominal rates) equal to 1 asymptotically, the relative importance of the various shocks and a small sample size may complicate this assessment. In order to characterize the estimated regression coefficient we perform some exercises varying the sample size (considering $T = 100, 200, 500$ thinking in quarterly data) and the standard deviation of the shocks to the target ($\sigma_{\varepsilon^{\overline{\pi}}}$). In the latter case we want to consider situations where the target changes little ($\sigma_{\varepsilon^{\overline{\pi}}}$ relatively small), values that result in a variance of the first differences of inflation close to that observed in the U.S. for a full post 1960 sample and some value in between, closer to the data in a post inflation targeting (or post 1984) sample. The top panel of Figure 1 displays the results, considering the empirical distribution (across 4000 samples of a given size) of the OLS estimator of a regression of inflation on nominal interest rates, the so-called levels estimator in a cointegration setting. It is quite evident that the distribution of the OLS estimator is never centered around 1 and it is often highly skewed to the left of 1. The bias can be quite substantial for the lowest value of $\sigma_{\varepsilon^{\overline{\pi}}}$ considered, as expected, since the (permanent) shifts in inflation and nominal rates are small, making it harder to identify the one-for-one relation. Naturally, as the sample size increases the distributions get more concentrated around 1, but we highlight that with the smaller $\sigma_{\varepsilon^{\overline{\pi}}}$ it takes many observations for the mass of the distribution to start concentrating around 1. There are several possible ways to reduce these biases. One option is to use filtered data. The bottom panel of Figure 1 displays the distribution of the OLS estimator when we pre-filter inflation and nominal interest rates using simple moving averages of length M, with M=40, or 10 years of data.[12] Clearly, the moving averages are effective in getting the estimates more concentrated around 1, and at a pace faster than what obtains with the raw data (levels) estimation. The speed of this convergence is obviously still dependent on the size of the sample and the relative importance of the shocks to the target. Filtering the data in the presence of cointegration can thus be a good option in empirical work to reduce these biases. However, we should stress that within the stationary world this procedure is quite useless, in the sense that it does not necessarily reveal any long-run relation, as McCallum (1984) pointed out long ago. In a stationary world the long run is a constant, and there is no way to identify a long-run relation by looking at two constants.[13]

12. The bias of the levels estimator is well-known, see Banerjee et al. (1993) and several alternatives are available to reduce this bias, accounting for short-run dynamics, see e.g. Saikkonen (1991) and Stock and Watson (1993). We perform this exercise just for illustrative purposes.

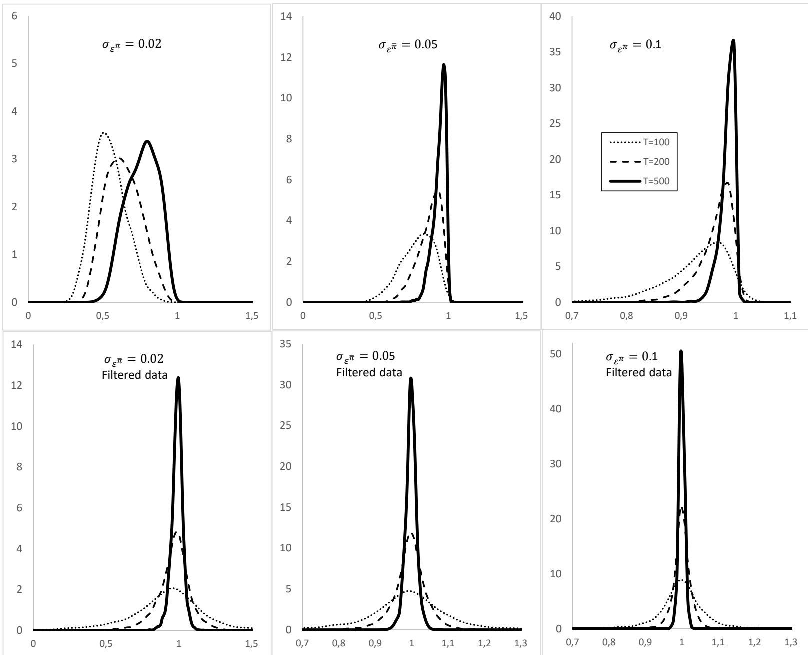
13. In this specific context, looking at the closed form solutions of the stationary model it can be seen that there is not much to be expected from taking moving averages of the data before proceeding to the regression. As an example take the simple NK model with just a monetary shock. The closed-form solution is $\pi_t = \frac{-\lambda\sigma}{\lambda\sigma\varphi_\pi+1}\varepsilon_t^m$ and $R_t = \frac{1}{\lambda\sigma\varphi_\pi+1}\varepsilon_t^m$ If we take moving

Within this model, it is also instructive to analyze the effects of shocks to the inflation target and typical monetary shocks (to the Taylor rule). One can interpret the inflation target shocks as permanent monetary shocks and Taylor rule shocks as temporary shocks. We will later attempt to identify empirically permanent and temporary monetary shocks, see Section 4. Figure 2 displays impulse response functions, the reactions of inflation, nominal rates and output to permanent ($\varepsilon_t^{\overline{\pi}}$) and temporary ($\varepsilon_t^m$) monetary shocks.

The effect of the temporary monetary shock is the standard one obtained in New Keynesian models: inflation and output fall temporarily after an unexpected increase in nominal rates. The effect of the permanent (inflation target) shock turns out to be temporarily expansionary. This occurs since not all firms reoptimize prices taking into account the new target for inflation. Hence, aggregate inflation is still below the new target and the central bank does not increase nominal rates as much as the new target would imply. The real interest rate thus falls and output expands. Also, this permanent shock leads obviously to permanently higher inflation and higher nominal rates. The level of the two variables changes by the same amount on account of the assumed cointegration relation with unitary coefficient. What we highlight is that the short-run response of inflation and nominal rates is similar to the long-run response, as both variables increase on impact. There would be a slight overshooting in the case of inflation if we considered persistence in the Taylor rule smoothing parameter ρ.[14]

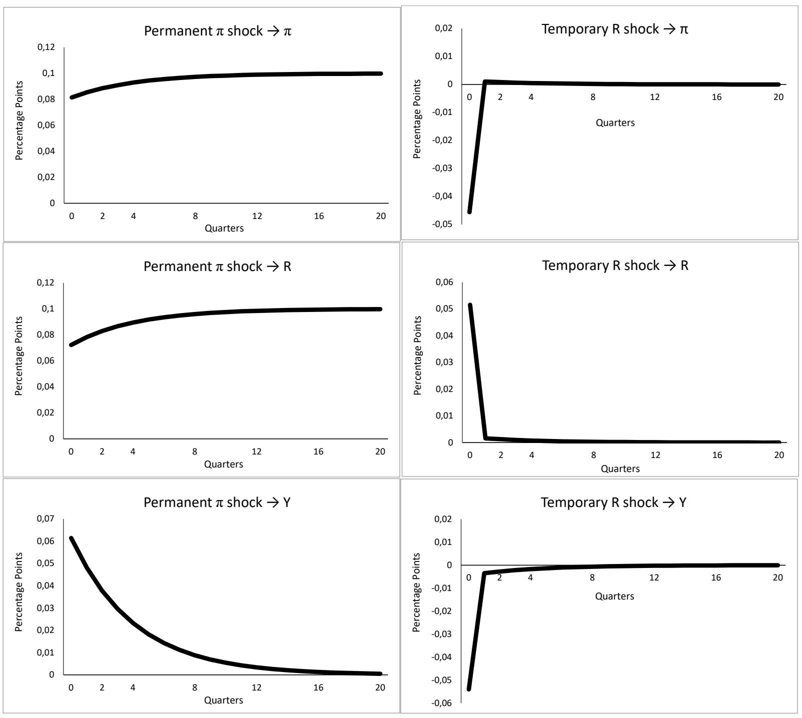
averages of length M of π_t and R_t (assuming no serial correlation in the shocks), we would still get the same theoretical OLS coefficient (not exactly empirically since M observations are lost). The characterization is not as simple in the presence of various persistent shocks but a coefficient moving towards 1 is not to be expected except in (very) special cases.

14. We have tried several parameterizations and this seems to be the crucial parameter affecting qualitatively the short-run responses of inflation, nominal rates and output.



Regression of inflation on nominal interest rates using data from the NK model with a varying target for inflation. Empirical Density Function of the OLS estimates across 4000 samples of simulated data with size $T = 100, 200, 500$, for different standard deviations of the shock to the inflation target: $\sigma_{\varepsilon^{\overline{\pi}}} = 0.02, 0.05$ or 0.1. Top Panel - Raw data; Bottom panel - filtered data with moving average of size M=40. M observations are disregarded and hence the effective sample sizes are $T = 60, 160, 460$.

Figure 1
Empirical Distribution of OLS estimates of Cointegration parameter - NK model with varying target



Response of inflation, nominal rates and output to permanent ($\varepsilon_t^{\overline{\pi}}$) and temporary ($\varepsilon_t^m$) monetary shocks. Standard parameterization. Deviations from the steady state in percentage points.

Figure 2
Impulse Response Functions of nominal shocks - NK model with varying target

3. Empirical strategy and results

3.1. Data

For our investigation, we have used annual data from 41 countries compiled by the OECD, drawing also on statistics of the IMF. Data pertains to the period ranging from 1961 to 2016, whenever available. CPI inflation was used, as well as short rates as defined by the OECD and IMF data on discount rates or monetary policy rates to extend the series backwards whenever OECD data (for the “substitute") T-bill rate was incomplete. More information on the data is provided in the appendix.

3.2. Cross-sectional analysis

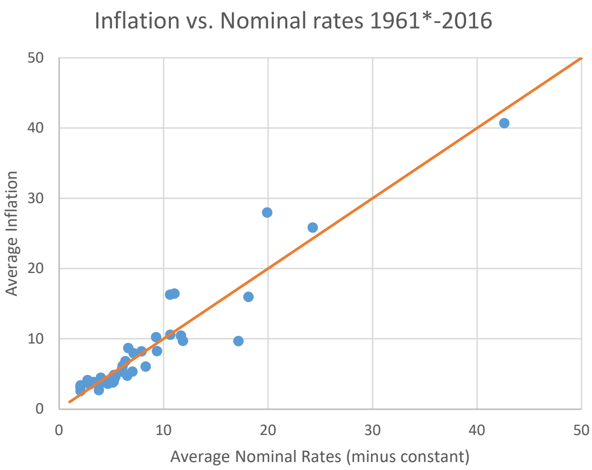
To analyze the relationship between inflation and nominal rates in the data, we want to examine a panel of countries $i = 1, \ldots, N$ for time periods $t = 1, \ldots, T$. Figure 3 plots average inflation against average nominal interest rates by country for the whole sample. The fit to a line with unitary slope is remarkably good. Next we plot the same data for two sub-samples in Figure 4, one from 1961 until the date of implicit or explicit adoption of inflation targeting (by country, and only if applicable) and another from that period until 2016. The specific break dates for inflation targeting (denoted IT henceforth) and justification can be found in the appendix. One can immediately conclude that the dispersion of inflation and nominal interest rates has decreased significantly in the most recent period. However, a quick look at the chart does not suggest any other difference in the relationship between the two periods. The same conclusion would obtain if we picked a common split sample date such as 2000 or 1995.

We also run a simple regression of the form:

$$\overline{\pi}_i = \alpha + \delta \overline{R}_i + \varepsilon_i \tag{10}$$

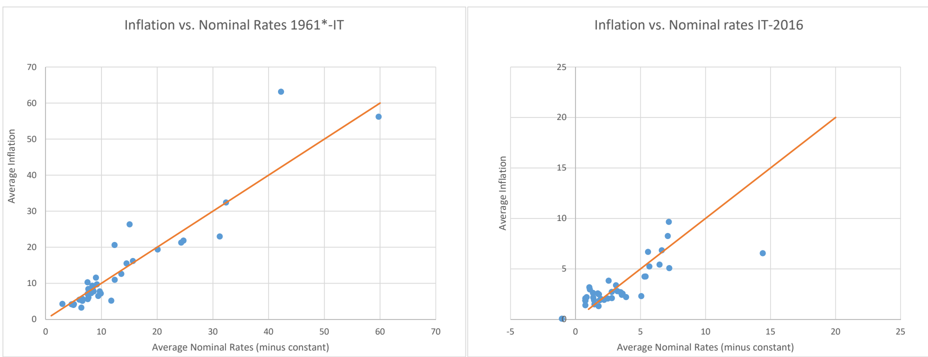
where bars denote the abovementioned averages across $t = 1, ..., T$ for each country i, and assess whether δ equals 1. If $\delta = 1$ then differences in average real rates across countries will be part of the error term. Meaningful OLS estimation of this common δ would require that the error term is ‘independent’ of average nominal rates, which could be a heroic assumption. E.g., average money growth (affecting average inflation) could move systematically with average nominal interest rates, even though it is conceivable that it does not. On the other hand, it is conceivable that countries with high nominal rates are those with unusually high real rates (low ε_i); this will bias our slope downwards.

Results are presented in Table 2. For the whole sample and for the first part of the sample the estimated slope is remarkably close to 1 and not statistically different from 1; the fit as measured by the R^2 is also remarkably high. One should notice that if country by country the relation were one-for-one, in the cross section this could still be difficult to identify, as countries can exhibit large disparities in average real rates. In this case, however, it seems the existence of several countries



Average CPI inflation plotted against average nominal rates (minus a constant) for the 41 countries. The constant is just the average difference between the two variables. For each country, the average of the variables is from 1961 (or later if data is unavailable) until 2016. The fit to the unit slope line is remarkable.

Figure 3
Inflation and nominal interest rates



Average CPI inflation plotted against average nominal rates (minus a constant). The constant is just the average difference between the two variables. Left Panel: for each country, the average of the variables is from 1961 (or later if data is unavailable) until the Inflation targeting break date. DE and CH are excluded. Right Panel: for each country, the average of the variables is from the Inflation targeting break date until 2016. 41 Countries included. The fit to the unit slope line is remarkable in the pre-IT sub-sample. In the post-IT period the relation between the two variables is not as clear but the residual dispersion around the unit slope line is lower than the one obtained in the Pre-IT sample (even lower than the one obtained in the full sample).

Figure 4
Inflation and nominal interest rates by country before and after inflation targeting

with high inflation and high nominal rates is enough to make unimportant even relatively large variations in real rates.

Next, for the inflation targeting period, we find a considerably lower coefficient, 0.57, and a worse fit as measured by the R^2. The same obtains with 2000 as a (common) break date. An obvious challenge to this approach is the little variation in the variables across countries. Most economies have been able to achieve low inflation rates, around a common target of 2%. There are still some differences in average nominal rates, likely stemming from differences in real rates, but the variability of inflation does not reveal a one-for-one relation. We further notice that in the post-IT (or post 2000) sample, only a few countries with high inflation and high nominal rates are helping in the identification of the positive slope; if we had focused on countries with average inflation below 5% the estimated slope would be even lower (cf. Figure 4). Now, this can well occur even if a common model featuring a Fisher equation generates this data. Actually, and very importantly, a Fisher relation seems consistent with the data: once we impose a slope coefficient equal to 1, the fit, as measured by the residual dispersion around this line is as low (even smaller) than the residual dispersion obtained for the pre-IT period (or for the whole sample), see the right columns of Table 2.

Period Benchmark	Free estimation				$\delta = 1$		
	δ	Const.	R^2	σ_ε	Const.	R^2	σ_ε
1961*-2016	0.99	-0.90	0.90	2.44	-0.99	0.90	2.44
	0.05	0.61			0.37		
Break Date is IT							
1961*-IT	1.07	-1.97	0.87	4.71	-1.03	0.87	4.77
	0.07	1.22			0.75		
IT-2016	0.57	0.60	0.63	1.24	-1.34	0.28	1.73
	0.07	0.37			0.27		
Break Date is 2000							
1961*-2000	1.05	-1.82	0.90	4.62	-1.07	0.65	4.67
	0.06	1.09			0.72		
2000-2016	0.63	0.54	0.85	1.19	-1.17	0.55	2.05
	0.04	0.27			0.32		

Regression of (averages of) inflation on constant and the nominal interest rate. Standard errors are below the estimates. R^2 is calculated as 1 minus (sum-of-squared of residuals divided by total variation in inflation). Averages of the variables are from 1961 (or whenever data is available) until 2016, denoted 1961*-2016, from 1961 (or whenever data is available) until IT adoption, denoted 1961*-IT and from IT adoption until 2016, denoted IT-2016. 2000 as break date is also considered. All Countries included for 1961*-2016, IT-2016, 1961*-2000 and 2000-2016 (41 observations). 1961*-IT does not include DE and CH. σ_ε is the standard deviation of the residuals.

Table 2. Regression Results

3.3. Panel cointegration

We use cointegration methods within a panel data approach to estimate the long-run relationship between inflation and nominal interest rates. For such an assessment an essential requirement is that both variables have a unit-root. Now, it is obviously possible that inflation and nominal rates are stationary in some countries (say, because of successful inflation targeting), non-stationary in other countries or even non-stationary during some period and stationary afterwards within the same country (e.g., if countries move from 1970's style stagflation policies to inflation targeting). We start by bypassing this issue, and consider the 41 countries in the regressions. The unit-root in both variables might be a reasonable hypothesis for almost all countries at least in the full sample covering 1961 (or whenever data for the two variables becomes available for each country) through to 2016.[15] The following panel cointegration model is considered:

$$\pi_{i,t} = \alpha_{0,i} + \delta R_{i,t} + u_{i,t}$$
$$\text{where } \Delta R_{i,t} = v_{it}$$

$w_{it} = (u_{it}, v_{it})'$ is possibly serially correlated but assumed independent across $i = 1, ..., n$, which is again heroic but only problematic for inference. $(1, -\delta)$ is a cointegrating vector and the equilibrium error $(\pi_{i,t} - \delta R_{i,t})$ is allowed to have country-specific fixed effects. If $\delta = 1$ this amounts to allowing for different real interest rates across countries. To estimate the model we consider several variations of the Dynamic OLS (DOLS) and Fully Modified OLS (FMOLS) panel estimators, see e.g., Phillips and Moon (1999), Pedroni (2000, 2001), Kao and Chiang (2000) and Mark and Sul (2003). We analyze the full sample 1961-2016 and also the Pre-IT and Post IT sub-samples. 2000 as a break point is also considered. We should underline that consideration of different break points across countries and data availability results in highly unbalanced panels, which makes inference problematic under more reasonable assumptions on the cross-sectional dependence of the data. We deal partially with this issue below. Table 3 contains the results.

For the full sample and for the earlier sub-samples (1960*-IT and 1960*-2000) we find a cointegration coefficient close to 1 in several instances, but lower values are also obtained. What is clear is the lower coefficient in the later part of the sample (Post IT or 2000-2016). This lower coefficient could be due to successful inflation targeting which likely results in stationarity of inflation and nominal rates in several countries over the more recent period. We should recall that for most countries (see Figure 4 but also Figure 5 in section 4.4 below) there are basically two inflation regimes throughout the sample: one running up until the 80s, characterized by high inflation, followed by a period of slow disinflation that ends with successful inflation targeting. By taking the whole sample or, in a lesser extent, a Pre-IT sample, we

15. Dickey-Fuller tests for each country confirm this idea, but we are surely aware of power considerations due to the short span of data in several countries.

Period Benchmark	Estimation of Cointegration parameter δ			
	DOLS		FMOLS	
Break Date is IT	Grouped	Pooled	Grouped	Pooled
1960-2016	0.76	1.08	0.86	0.95
	0.03	0.02	0.03	0.01
1960*-IT	0.76	0.94	0.72	0.77
	0.06	0.04	0.06	0.02
IT-2016	0.40	0.35	0.38	0.38
	0.08	0.03	0.07	0.02
Break Date is 2000				
1960*-2000	0.99	0.57	0.77	1.00
	0.05	0.05	0.08	0.02
2000-2016	0.44	0.45	0.51	0.49
	0.07	0.02	0.06	0.03

Estimation by DOLS and FMOLS as implemented in Eviews considering country fixed-effects in the cointegration equation. We consider Grouped and Pooled (weighted) estimation. Pooled (weighted) estimation considers cross-section estimates of the error covariances. With FMOLS we consider heterogeneous first-stage long-run coefficients. Using DOLS the u_{it} are allowed to be correlated with at most p_i leads and lags of v_{it}. We choose p_i according to the AIC criterion. HAC standard errors are below the estimates. For each country, the sample runs from from 1961 (or whenever data is available) until 2016, denoted 1961*-2016, from 1961 (or whenever data is available) until IT adoption, denoted 1961*-IT and from IT adoption until 2016, denoted IT-2016. 2000 as break date is also considered. All Countries included for 1961*-2016, IT-2016, 1961*-2000 and 2000-2016 (41 observations). 1961*-IT does not include DE and CH.

Table 3. Panel Cointegration Regression Results

are able to exploit these permanent shifts in inflation and nominal rates. The time series evidence below will reinforce this idea.

Next we try to be more precise in the inferences made, at the cost of losing several countries, since we require a balanced panel. We begin by testing for non-stationarity of nominal rates and inflation in our dataset using the method developed in Smith et al. (2004), which deals explicitly with cross-sectional dependence that most likely characterizes our data.[16] We use 4 specific tests statistics described in Smith et al. (2004): $\bar{t}$, $\overline{LM}$, $\overline{max}$ and $\overline{min}$ with the null of a unit-root and alternative hypothesis of stationarity of at least one of the series, including a constant but no trend. Due to data availability we apply the unit root tests on inflation to a panel of 26 economies from 1962 to 2016 and also to a panel of 31 economies from 1974 to 2016. For the nominal interest rate we have a

16. We avoid employing the tests due to Levin, Lin and Chu (2002), Breitung (2000), Im, Pesaran and Shin (2003), Maddala and Wu (1999), Choi (2001) and Hadri (2000) as they do not consider this cross-sectional dependence.

panel of 24 economies from 1968 to 2016 and a panel of 27 economies from 1974 to 2016.[17] Results are presented in Table 4.

Variable	Inflation		Interest rate	
Test	P-values			
$\bar{t}$	0.206	0.000	0.952	0.921
$\overline{LM}$	0.740	0.000	1.000	0.743
$\overline{max}$	0.005	0.999	0.670	0.642
$\overline{min}$	0.125	0.036	0.999	0.681
N	26	31	23	26
t_0	1962	1974	1968	1974

In the four tests we use bootstrapping methods to account for potential cross-sectional dependence in the data. We start by estimating Augmented Dickey-Fuller equations for each country, with the number of lags chosen using AIC . We construct bootstrap samples by estimating a new ADF equation through OLS, using the number of lags chosen initially for each country, but this time we impose a unit root. Residuals from this estimation are stored and will be later used for bootstrap sampling. Following Stine (1987) these residuals are recentred. Resampling is used to generate bootstrap innovations from the vectors of recentred residuals, ensuring that the cross-sectional dependence in the data is maintained. With each sample of innovations, a new panel is constructed with an imposed unit-root. We construct 5000 bootstrap panels in such way and then apply the tests to it to compute p-values that are robust to the specific cross-sectional dependence present in the data. Common set of countries in all tests : Australia, Austria, Belgium Canada, Denmark, Finland, Germany, Iceland, India, Ireland, Italy, Japan, Korea, The Netherlands, Norway, Portugal, South Africa, Spain, Sweden, Switzerland, United Kingdom, United States. For inflation, when N=26 France, Greece, New Zealand and Turkey are included; when N=31 Chile, Colombia, Indonesia, Israel and Mexico are additionally included. For the nominal interest rate, when N=23 Colombia is included; when N=26 France, Greece and New Zealand are additionally included.

Table 4. Panel unit root tests with cross-sectional dependence

As found in the literature, the unit-root in nominal interest rates seems to be quite evident whereas for inflation the results are less conclusive. We proceed by inspecting results of individual Dickey-Fuller tests, in search of the countries likely contributing to rejection of the null and prune our dataset accordingly. We arrive at a restricted panel of 20 countries with (common) data starting in 1963 for which we cannot reject the null of a unit-root at the 1% level in any of the tests. We then estimate and test for cointegration in this restricted dataset.[18] Table 5 presents

17. The Smith et al. (2004) tests, because they rely on a bootstrap of a vector of estimated residuals, require a balanced panel, thereby restricting the number of countries in the panel.

18. As previously discussed, inflation targeting may have changed the nature of these variables from non-stationary to stationary. Trying to assess this could be interesting, but such an assessment is hard to make given the constraints imposed by the availability of data. Once we divide our sample in 2 periods, the power of our test is reduced significantly, as can be seen in the tables presented in Smith et al. (2004). Even for a sample with no cross-section dependence, if there is serial correlation such that lags must be used in the augmented Dickey-Fuller regression, the power of the test $\overline{max}$ (the most powerful one) with $\rho = 0.9$ and 4 lags goes from 0.983 when $T = 50$ and $N = 25$, to 0.451 when $T = 25$ and $N = 25$.

the estimation results, considering also a particular case of the cointegration model above, where $\alpha_{0,i} = \alpha_0$ (or no fixed effects in the cointegration relation). If $\delta = 1$ this imposes the same long-run real interest rate across countries, which could well be reasonable given the nature of the restricted panel (comprised of developed and quite open economies).

	Estimation of Cointegration parameter δ			
Fixed Effects	DOLS		FMOLS	
Period	Grouped	Pooled	Grouped	Pooled
1963-2016	0.66	0.63	0.71	0.59
	0.04	0.05	0.05	0.01
1963-2000	0.63	0.61	0.70	0.43
	0.07	0.07	0.08	0.02
2000-2016	0.33	0.32	0.37	0.46
	0.07	0.04	0.06	0.04
No Fixed Effects	DOLS		FMOLS	
Period	Grouped	Pooled	Grouped	Pooled
1963-2016	0.78	0.76	0.82	0.76
	0.02	0.03	0.03	0.01
1963-2000	0.79	0.77	0.81	0.75
	0.03	0.04	0.03	0.01
2000-2016	0.64	0.63	0.65	0.77
	0.04	0.01	0.04	0.02

Estimation by DOLS and FMOLS as implemented in Eviews considering country fixed-effects and no fixed effects in the cointegration equation. We consider Grouped and Pooled (weighted) estimation. Pooled (weighted) estimation considers cross-section estimates of the error covariances. With FMOLS we consider heterogeneous first-stage long-run coefficients. Using DOLS the u_{it} are allowed to be correlated with at most p_i leads and lags of v_{it}. We choose p_i according to the AIC criterion. HAC standard errors are below the estimates. For each country, the sample runs from 1963 until 2016, from 1963 until 2000, and from 2000 until 2016. 20 Countries included: Austria, Belgium Canada, Denmark, Finland, Germany, Iceland, Ireland, Italy, Japan, Korea, The Netherlands, Norway, Portugal, South Africa, Spain, Sweden, Switzerland, United Kingdom, United States.

Table 5. Panel Cointegration Regression Results, 20 Countries

For the full sample and for the earlier sub-sample (1963-2000) we find a cointegration coefficient clearly below one if fixed effects are permitted and hovering around 0.8 otherwise. In both instances, it is clear that a lower coefficient obtains in the later part of the sample. Again, this lower coefficient could be due to successful inflation targeting resulting in stationarity of inflation and nominal rates in several countries over the more recent period. The coefficient below 1 within the more clearly non-stationary period is more troublesome. It can be the result of the slow fall in real rates that has been documented in developed economies, see, e.g., Holston, Laubach and Williams (2017) and the discussion in Gordon (2014) or

Summers (2014). In any case, it could still be that the data is consistent with $\delta = 1$ on the grounds of (lack of) efficiency or biases in the estimation (see Westerlund 2008) due to the small number of cross-sections, the lack of control for cross-sectional dependence or even successful inflation targeting at low common values during a large part of the sample. We thus test for cointegration assuming $\delta = 1$, which means that we simply test for a unit-root in the (ex-post) real interest rate (allowing for differences in long-run real rates across countries). Table 6 shows that in the restricted sample we use, inflation and the nominal rate are likely characterized by a unit-root, as referred above, while the null of a unit-root in the real interest rate is clearly rejected in two of the tests and rejected at a 5% significance level in another one.[19]

Unit-Root Tests P-Values			
Test	Inflation	Nominal rate	Real rate
$\bar{t}$	0.371	0.825	0.021
$\overline{LM}$	0.897	0.998	0.255
$\overline{max}$	0.023	0.260	0.000
$\overline{min}$	0.418	0.937	0.005
N	20	20	20
t_0	1963	1963	1963

In the four tests we use bootstrapping methods to account for potential cross-sectional dependence in the data, see the previous table for details. 20 Countries included: the same as in the previous table

Table 6. Cointegration test with cross-sectional dependence

The results presented in Table 6, which complement the results found in Westerlund (2008), are evidence of a positive one-for-one relationship in nominal interest rates and inflation. Jensen (2009) argues that the difficulty in finding the Fisher relation in the data is simply a direct consequence of the lack of permanent shifts in inflation but our results suggest otherwise.

19. Testing for the Fisher relationship through cointegration has been done before in the literature. Most of the literature does not find cointegration, or fails to find it as a one-for-one relationship, see Mishkin (1992) or Crowder and Hoffman (1996). More recently, however, Westerlund (2008) shows how this inability to reject the null hypothesis of cointegration with unit coefficient may be due to the low power of time-series residual-based tests of cointegration, such as those due to Pedroni (1999), Pedroni (2004), Kao (1999) and Maddala and Wu (1999). Using a panel of quarterly data from 20 OECD countries that spans from 1980 to 2004, that paper shows that the Fisher relation cannot be rejected when one looks at the panel evidence. Our test differs from Westerlund's in that we do not model cross-sectional dependence of real interest rates through common factors (with the cost of losing some power relatively to his approach if that parsimonious structure is valid).

3.4. Time-Series evidence

In this section we analyze time series evidence on the relation between nominal rates and inflation. We could just repeat the cointegration analysis above country by country. If inflation and nominal rates are integrated of order one and cointegrated, one can just run the following regression:

$$\pi_t = \alpha + \delta R_t + \varepsilon_t \tag{11}$$

where ε_t is allowed to be serially correlated. This would deliver a unitary δ if the Fisher relation holds in the long run. Due to small sample biases, employing moving averages can enhance efficiency in the estimation of δ as short-run fluctuations are averaged out. Furthermore, in this case one could expect an improvement in the fit to the line with unit slope, measured by an increase in the R^2. If, on the contrary, inflation and nominal rates are stationary running this regression is obviously quite unadvised as the Fisher relation is just one of the equations that determines the evolution of the economic system. Monetary policy, represented here by R_t, likely responds to changes in ε_t, which could be changes in some natural real rate, in money growth or in the economy's slack. Again, employing moving averages in this case is of no help as McCallum (1984) stressed long ago. We go back to a common theme in this paper, that permanent shifts in the variables are needed to meaningfully identify the relation using reduced form methods, i.e. without requiring a complete knowledge of the economic system, as King and Watson (1997) recalled. But again, filtering the data can nonetheless be helpful in uncovering the long-run relation in a non-stationary environment, as the analysis in Section 2.1 suggested.

Table 7 presents the results of running regressions of inflation on interest rates using raw data and moving averages of length $M = 2, 5, 10$ years of data. We run regressions for each country separately and present the average estimate of δ and the standard deviation of the estimates across countries. We report also the average R^2 of these regressions and the average standard deviation of the residuals. We also report the average standard deviation of the residuals computed upon imposing $\delta = 1$. We look at the whole sample and also at pre-IT and post- IT samples using the break dates previously discussed.

For the sample as a whole and for the pre-IT sample the estimated slope is around 0.80 on average, using the raw data ($M = 1$) or moving averages of length $M = 2, 5, 10$, but with more dispersion across countries in the first part of the sample and a somewhat lower value for $M = 10$ in this case. The fit, as measured by the R^2, is relatively high in these samples, but higher when considering the whole sample. The residual dispersion is correspondingly lower for the whole sample and naturally decreasing in M as more noise is filtered out. Also, imposing a unitary slope line leads to a relatively small increase in the residual dispersion. Turning to the post-IT sample one observes a clear fall in the average estimated slope to values around 0.40 for every M and some variation across countries (comparable to the pre-IT sample). The R^2 also decreases somewhat compared to the full sample.

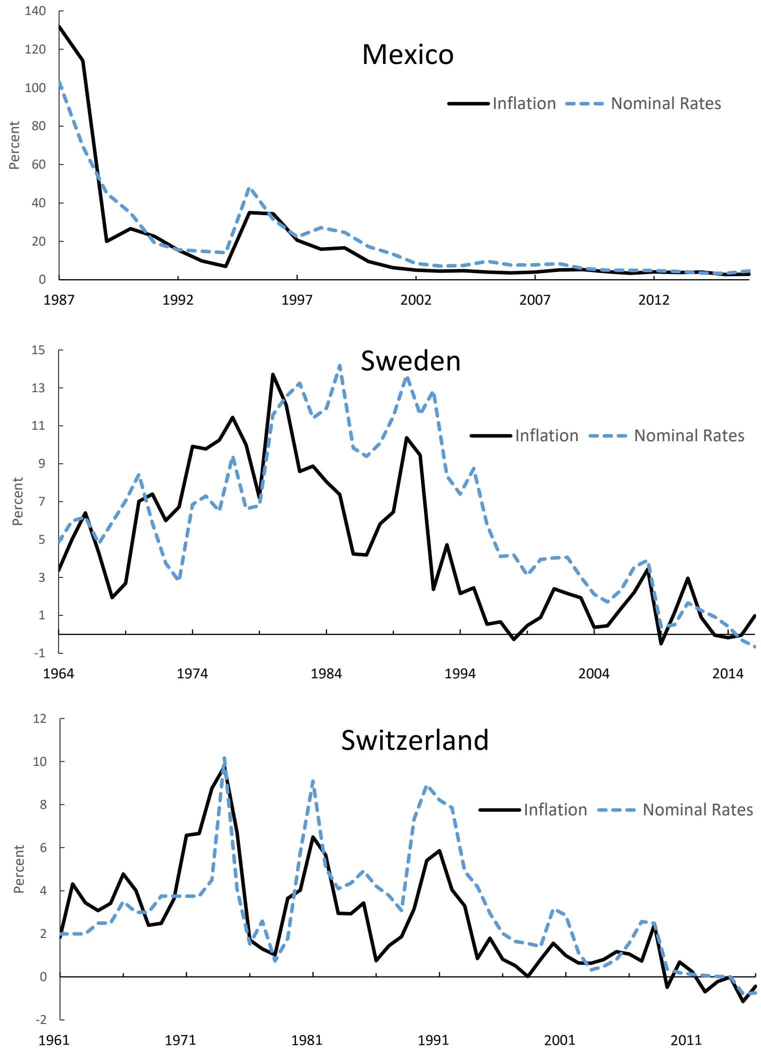
Very importantly, and similarly to the cross-sectional evidence, once we impose a slope coefficient equal to 1, the average residual dispersion around this line is rather small compared to the residual dispersion obtained for the pre-inflation targeting period (or for the whole sample), even in the case of an estimated slope (lower than 1 as it turns out).

We interpret these results as follows. When we look at the whole sample we observe permanent shifts in inflation and nominal rates in most countries; an initial phase with high inflation (and high nominal rates) and the later part with low inflation (and low nominal rates). For the pre-IT sample there is also strong variation across time in the two variables, within a high(er) inflation environment. This permits the identification of a strong positive relation between inflation and nominal rates in both samples, although the estimated coefficient is lower than one, which could again indicate a fall in real interest rates across time. The later part of the sample is characterized by little variability in both inflation and nominal rates; the two variables could well be characterized as stationary in most countries, which results in a much lower average regression coefficient. Figure 5 illustrates this point graphically. Again, this data can be seen as consistent with a Fisher relation, in view of the very low observed (average) residual dispersion around the line with unitary slope.

Period Benchmark		M			
1961*-2016		1	2	5	10
	Avg. $\hat{\delta}$	0.80	0.80	0.80	0.81
	Stdv $\hat{\delta}$	0.13	0.11	0.11	0.10
	R^2	0.54	0.57	0.62	0.64
	σ_ε	4.64	3.85	2.87	2.27
	σ_ε $(\delta = 1)$	5.74	4.82	3.57	2.78
1961*-IT					
	Avg. $\hat{\delta}$	0.79	0.81	0.73	0.57
	Stdv $\hat{\delta}$	0.29	0.26	0.22	0.19
	R^2	0.37	0.38	0.42	0.37
	σ_ε	6.04	5.00	3.58	2.81
	σ_ε $(\delta = 1)$	8.16	6.87	4.93	4.13
IT-2016					
	Avg. $\hat{\delta}$	0.43	0.40	0.34	0.43
	Stdv $\hat{\delta}$	0.20	0.19	0.22	0.22
	R^2	0.37	0.39	0.44	0.55
	σ_ε	1.45	1.15	0.65	0.27
	σ_ε $(\delta = 1)$	2.24	1.95	1.39	0.83

Time-series regressions of inflation on nominal rates for the 41 countries, subject to data availability. We report the average OLS estimate across countries, the standard deviation of the OLS estimates, the average R^2 of the regressions, the average standard deviation of the residuals and the average standard deviation of the residuals considering a unitary slope. The whole sample and the Post-IT sample include all countries except Turkey and Indonesia when $M = 10$ due to insufficient data. The pre-IT sample does not include DE and CH and it does not include, due to insufficient data, and at least for some $M > 1$, Chile, Estonia, Slovak Rep. Slovenia, Brazil, China and Russia. Results excluding always these countries are very similar.

Table 7. Time-series regressions of inflation on nominal rates. Annual data and moving averages of length $M = 2, 5, 10$



Mexico has periods with very high and very low inflation, the regression coefficient is close to 1; Sweden also has periods with high and low inflation, but real rate seems to fall since the mid 1990s; the estimated coefficient is around 0.7. Switzerland is characterized by relatively low inflation and nominal rates throughout the sample, making the relation harder to identify even if there is comovement; the estimated coefficient is also around 0.7. In all three cases, if we focus on the later part of the sample, characterized by low inflation, low nominal rates, and little dispersion in the two variables, the regression coefficients fall substantially.

Figure 5
Data - Inflation and Nominal Rates

4. Long-run and short-run dynamics in the U.S., Japan and France

We now move from the characterization of the long-run relation between inflation and nominal rates to assessing the effects of identified nominal shocks on inflation and nominal rates. In view of the analysis above, we aim at distinguishing clearly between permanent and transitory nominal shocks. Permanent shocks ought to drive the long-run positive relation between inflation and nominal rates, since the two variables appear to be cointegrated whenever they are better characterized by a unit-root. Further, it is important to understand how long it takes for the adjustment to take place. Next, such permanent shocks may coexist with temporary, or mean reverting, nominal shocks, and it may well be the case that a temporary shock leading to a rise in nominal interest rates eventually pushes inflation down, as is typically concluded in the literature, see the comprehensive review of Christiano, Eichenbaum and Evans (1999), the references therein, and the extensive literature that followed.

We take monthly data for the levels of core inflation, nominal interest rates (see Figure 6), and industrial production (our measure of output) for the U.S. (1961-2018), Japan (1961-2018) and France (1961-2018) and investigate the dynamics of these variables using a very simple structural vector error correction model (VECM), see the detailed description of the data in the appendix. The Supplemental Material appendix contains results for the U.K., Germany and the euro area.[20] We will assume inflation and nominal rates are cointegrated (the coefficient does not need to be one) and we impose that output is not cointegrated with those two variables. We then track the effects of identified permanent and transitory nominal shocks. Standard tests reveal that cointegration between inflation and nominal rates is a reasonable hypothesis if one takes the whole sample (and the two variables appear to have a unit-root, even in the case of the euro area). A unit-root also characterizes output. Assuming a vector autoregression representation for the data, if the three variables are integrated of order one and cointegrated they can be represented in an error correction framework. The VECM reduced form representation is given by:

$$\Delta X_t = \alpha_0 + +\gamma\beta X_{t-1} + \sum_{j=1}^{K_r} \beta_j \Delta X_{t-j} + u_t$$

$X_t = (\pi_t, R_t, Y_t)'$ collects inflation, nominal rates and output and $\gamma := (\gamma_\pi, \gamma_R, \gamma_Y)'$ collects the adjustment coefficients. $\beta := (1 : -\delta : 0)$ is the cointegrating vector, where we make explicit that output is not part of the cointegration relation. $u_t = (u_{\pi,t}, u_{R,t}, u_{Y,t})'$ is a vector of reduced form innovations. In what follows we follow very closely Lütkepohl (2006, Chapter 9). For identification of so-called structural shocks we assume these reduced form shocks are related to the structural shocks through some non-singular matrix B such

20. We chose to include here results for France as they are close to those obtained for the euro area in the period 2001-2018 while allowing us to investigate a longer sample.

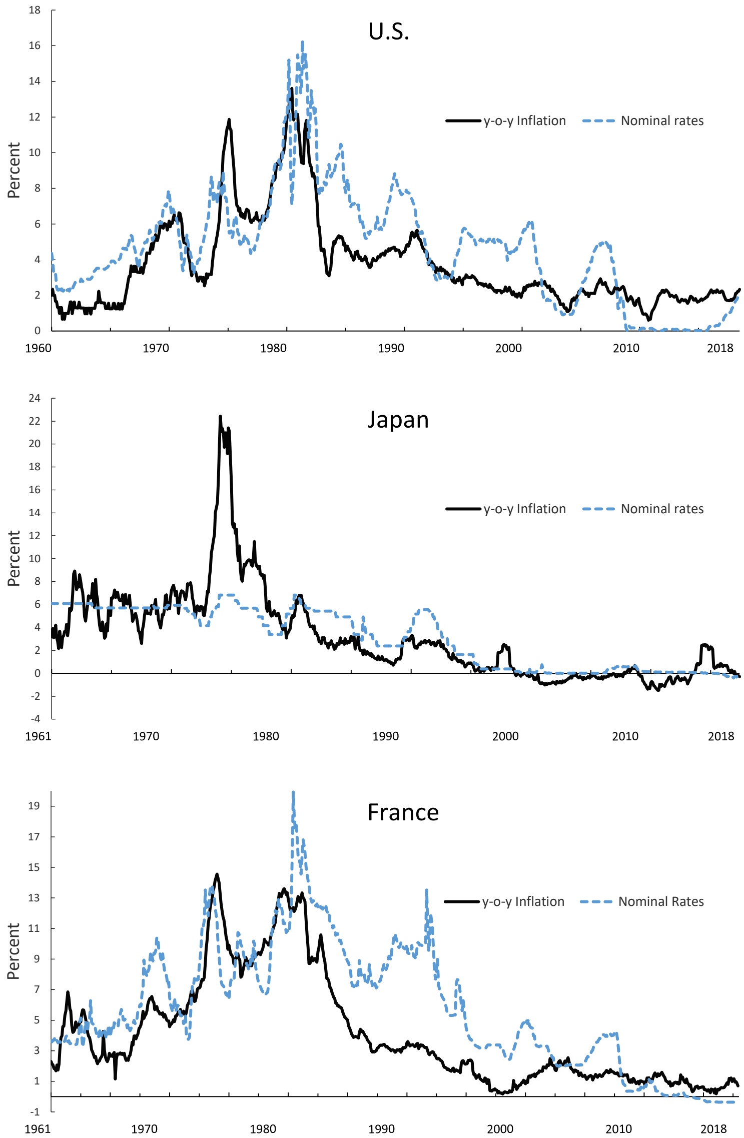


Figure 6
Data - Inflation and Nominal Rates - monthly data for U.S., Japan and France

that $u_t = B\varepsilon_t$, where $\varepsilon_t = (\varepsilon_{\pi,t}, \varepsilon_{R,t}, \varepsilon_{Y,t})'$ is a vector of serially and mutually uncorrelated structural shocks with normalized (unit) variance. Restrictions on B are needed to identify the structural shocks and assess their effects. Impulse response functions on the reduced form model are obviously meaningless. In view of the cointegration relation, it follows that this reduced form implies the following Beveridge-Nelson decomposition:

$$X_t = X_{0,t}^* + \Xi \sum_{i=1}^{t} u_i + \sum_{j=0}^{\infty} \Xi_j^* u_{t-j} \tag{12}$$

where the last term is absolutely summable and thus stationary. $X_{0,t}^*$ contains the initial values and possibly linear trends while the second term is the so-called stochastic trend. In structural form we have

$$X_t = X_{0,t}^* + \Xi B \sum_{i=1}^{t} \varepsilon_i + \sum_{j=0}^{\infty} \Xi_j^* B \varepsilon_{t-j} \tag{13}$$

Matrix Ξ is a straightforward function of the parameters of the reduced form and it has reduced rank if there is cointegration. We are assuming a cointegration rank of one so Ξ has rank 2. This facilitates the identification of B. Since B is a non-singular matrix then ΞB will have rank 2 as all the variables in X_t are integrated but π_t and R_t are cointegrated, thus sharing a stochastic trend. As a first step to identify B one can set one of the columns of ΞB to zero. This simply amounts to assuming that there is a temporary shock, i.e. one without permanent effects. This column should not be the third one (the loadings related to $\sum_{i=1}^{t} \varepsilon_{Y,i}$) since output should retain a stochastic trend, which is not, by assumption, the stochastic trend shared by π_t and R_t. Setting the first or second column of ΞB to zero amounts to establishing which of the shocks $\varepsilon_{\pi,t}$ or $\varepsilon_{R,t}$ is the temporary (as opposed to the permanent) nominal shock. Here we will pick $\varepsilon_{\pi,t}$ as the permanent shock and $\varepsilon_{R,t}$ as the transitory shock; the second column of ΞB is set to zero. Thus, only $\varepsilon_{\pi,t}$ will generate the common stochastic trend and hence have a permanent effect on π_t and R_t.[21] We need one extra restriction to identify the permanent shock to output, or locally identify B. It suffices to place restrictions on ΞB. We assume that the permanent output shock does not affect the long-run level of both inflation and nominal rates (the first two elements in the third column of ΞB are set to zero, but restricting only one of these to zero would result in an equivalent identification, since the other entry would be zero). Finally, we overidentify the model and assume that the permanent nominal shock does not affect output in the long run (i.e., we postulate long-run neutrality with regards to output) as common in the literature,

21. Given that the permanent nominal shock only has permanent effects on inflation and nominal rates, this choice is innocuous. We could alternatively pick $\varepsilon_{R,t}$ as the permanent shock and $\varepsilon_{\pi,t}$ as the transitory shock; the first column of ΞB would be set to zero such that only $\varepsilon_{R,t}$ would have a permanent effect on π_t and R_t. In this way we would obtain exactly the same impulse response functions as those obtained with the alternative identification, only the labeling of the shocks would be switched ($\varepsilon_{R,t}$ would be the permanent shock).

see e.g., Shapiro and Watson (1988), King, Plosser, Stock, and Watson (1991) or Galí (1992).[22]

Table 8 reports the estimation results regarding the cointegration vector and the adjustment parameters for the U.S., Japan and France, taking the whole sample and also the pre-IT and post-IT sub-samples, where the break date is 1984 for the U.S., 2001 for Japan and 1999 for France.[23] In the case of the U.S. we note a cointegrating parameter below one for the whole sample (0.63) a somewhat larger value for the pre-IT period (0.73) and a substantially lower value for the post-IT period, in line with the evidence presented above for several countries. For Japan the coefficient is somewhat larger than one for the whole sample and for the pre-IT sample but becomes negative (and hardly significant) in the post-IT sub-sample. This is certainly the result of the stationary behavior of inflation and nominal rates since 2001 in Japan; the two variables hover around zero. For France the cointegration coefficient is high over the whole sample and the pre-IT period but rather low albeit significant over the post-IT period. We notice further that the variable adjusting more significantly towards the cointegration relation is inflation (larger and significant coefficient in most instances, although for France nominal rates seem to adjust in relevant ways).

Turning now to the impulse response functions of the identified structural shocks for the U.S., Japan and France (Figures 7, 8 and 9) we note the following. For the three economies, the permanent output shock is associated with temporarily lower inflation and not very significant effects on nominal rates. For France the effects are hardly significant also for inflation. The effects of the temporary nominal shock are the conventional ones: a temporary shock increases nominal rates and drives inflation downwards. Output falls on impact in the U.S., although it recovers in less than one year. In the case of Japan the fall in output is more persistent whereas for the France the contractionary effects are small and hardly significant. All these impacts essentially fade away after three years, closer to four years in the case of Japan. We also notice that in the case of Japan the moderate increase in nominal rates is associated with a strong fall in inflation; for the U.S and France the effects are quantitatively more standard. Turning now to the permanent nominal shock (leading to a permanent increase in nominal rates) we notice that the effects on output are somewhat expansionary (not for France on impact, but here the effects are hardly significant). As expected, the permanent increase in nominal rates is accompanied by an increase in inflation towards the new long-run level. The ratio of the long-run levels reported for the two variables is obviously the

22. Empirically, not adding this extra assumption changes little the results since the estimated entry in ΞB is very close to zero. Alternatively, one could only set this restriction and allow for permanent effects of output shocks on inflation and nominal rates. We should notice that this would only affect somewhat the impulse response functions associated with the output shock, not the ones associated with the nominal shocks. Results are available upon request.

23. All the results in this section and in the appendix regarding estimation of the structural VECM model are obtained using JMulti, see http://jmulti.de/

	Estimation of Cointegration parameters δ, $\gamma := (\gamma_\pi, \gamma_R, \gamma_Y)$											
Sample	U.S.				Japan				France			
	δ	γ_π	γ_R	γ_Y	δ	γ_π	γ_R	γ_Y	δ	γ_π	γ_R	γ_Y
1961*-2018	0,63	-0,02	0,02	0,00	1,49	-0,02	0,004	-0,001	0,80	-0,008	0,02	0,00
	0,09	0,01	0,01	0,001	0,24	0,01	0,003	0,0005	0,16	0,005	0,01	0,000
1961*-IT	0,73	-0,05	0,005	0,00	1,77	-0,02	0,006	0.000	1,19	-0,006	0,03	0,00
	0,14	0,01	0,02	0,001	0,40	0,01	0,003	0.000	0,24	0,006	0,01	0,000
IT-2018	0,32	-0,05	-0.03	-0,002	-1,15	-0,05	-0,004	-0,001	0,23	-0,08	-0.025	0,00
	0,04	0,02	0,02	0,001	1,50	0,02	0,005	0,002	0,07	0,02	0,02	0,002

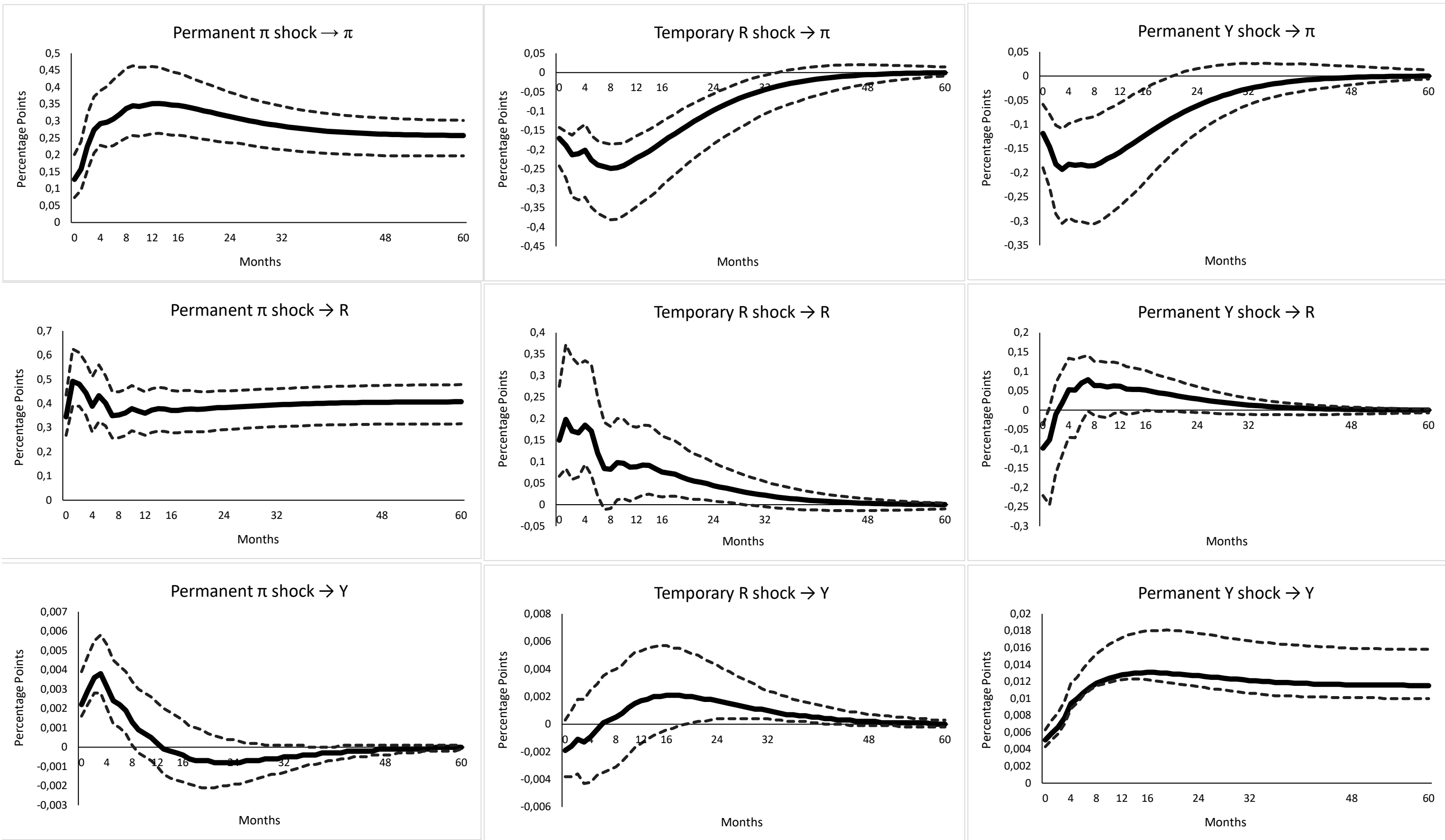
Estimates of cointegration parameters in the VECM model. Standard errors are below the estimates. Cointegration rank is restricted to 1 and coefficient of output in the cointegration relation is restricted to zero. 6 lags of differenced endogenous variables are considered in the VECM. Simple two step estimator (S2S), as implemented in JMulTi is employed. Estimation samples for the U.S: i) 1961*-2018, or 1961 M1 - 2018 M5 (T=689), ii) 1961*-IT, or 1961 M1 -1984M1 (T=282) and iii) IT-2018, or 1984 M1 -2018 M5 (T=413); Estimation samples for Japan: i) 1961*-2018, or 1960 M8 - 2017 M6 (T=683), ii) 1961*-IT, or 1960 M8 - 2001 M12 (T=497) and iii) IT-2018, or 2001 M1 -2017 M6 (T=198); Estimation samples for France: 1961*-2018, or 1961 M8 - 2018 M12 (T=689), ii) 1961*-IT, or 1961 M8 - 1999 M12 (T=461) and iii) IT-2018, or 1999 M1 -2018 M12 (T=240).

Table 8. Vector Error Correction model Estimation. Cointegration parameters

cointegration parameter. What we highlight is that this adjustment takes place in roughly less than two years in all cases while qualitatively the short-run effect of this permanent shock is very similar to the long-run one. No fall of inflation in the short run is observed, i.e., we also find the so-called Neo-Fisher effect first documented by Uribe (2017, 2018) in a somewhat different setting. In his setup the unitary cointegration coefficient between inflation and nominal rates is imposed whereas we estimate this parameter. Second, Uribe (2017, 2018) identifies temporary and permanent monetary shocks by restricting the impact response of inflation and output to a temporary monetary shock to be non-negative (i.e., the conventional effect is assumed for identification, at least on impact). We do not require that assumption. Despite these differences, qualitatively the results are very similar. Finally, we highlight that results in the appendix for the U.K., Germany and the euro area are strikingly similar to the ones reported here.

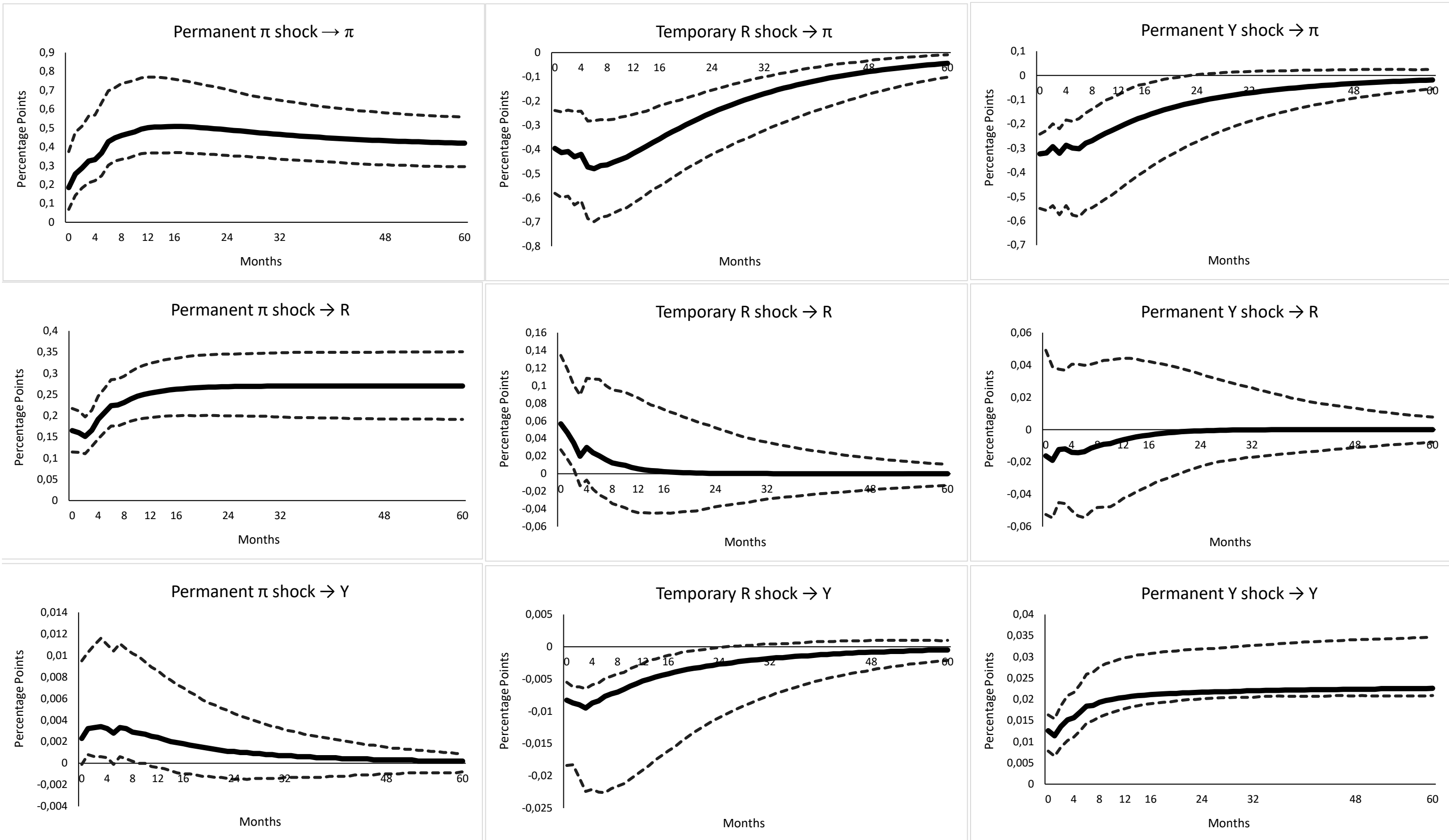
Next we analyze the decomposition of the variance of the forecast error for π_t, R_t and Y_t. Table 9 displays the results for the three economies. First, the bulk of the variance of the forecast error in the case of output is attributed to the output shock, and very clearly so at horizons greater than one year. This shock also accounts for around $25\% - 35\%$ of the variance of the forecast error of inflation in the case of the U.S. and Japan at short horizons and a very low percentage in the case of nominal interest rates. In the case of France this shock actually contributes very little to the variance of the forecast error of both inflation and nominal rates. Next, we highlight that the contribution of the permanent nominal shock for the variance of the forecast error of nominal rates is quite high even at short horizons (less so for France), ranging between 39% and 98% at horizons below one year.

This shock also accounts for an important fraction of the variance of the forecast error of inflation, with starting values at 27%, 11% and 72%, respectively, for the U.S., Japan and France while reaching 51% in the U.S. at the one year horizon, 49% in Japan at horizons greater than two years and much higher values in the case of France. The temporary nominal shock contributes most to the variance of the forecast error of inflation at short horizons in the U.S. and Japan, starting at around 50% in the case of the U.S. and Japan (and somewhat less, 26%, in the case of France) and then its contribution gradually diminishes with the forecast horizon. This temporary shock contributes little to the variance of the forecast error of nominal rates in the U.S. and Japan, with initial values around 10% (and gradually decreasing with the horizon). For France the contribution is above 50% up to one year horizons. Finally, the contribution of this temporary nominal shock to the variance of the forecast error of output at short horizons is somewhat relevant for the U.S. and specially Japan (10% and 29%, respectively at one quarter ahead), whereas for France the contribution is very small.



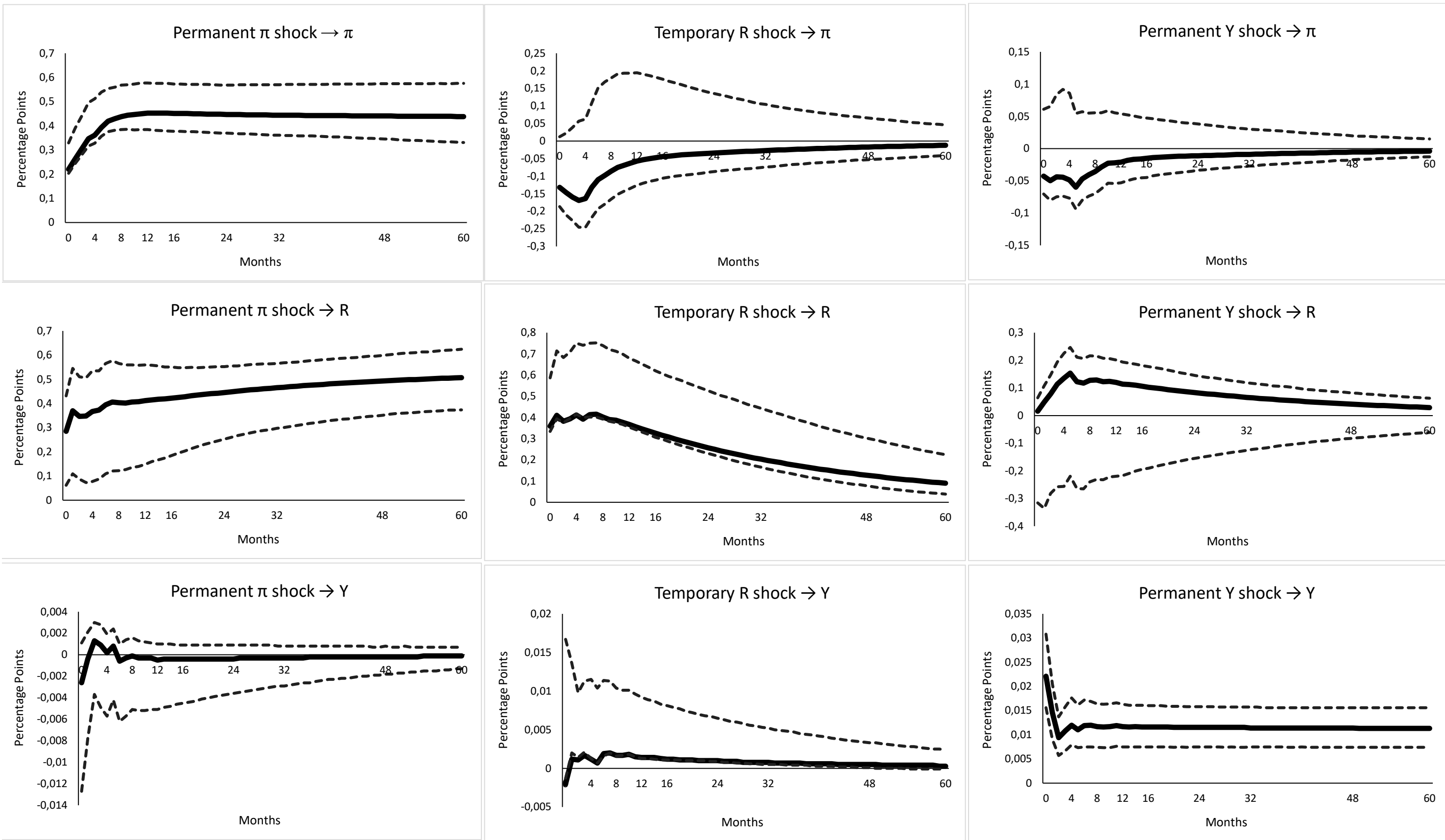
Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 7
Impulse Response Functions - U.S.



Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 8
Impulse Response Functions - Japan



Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 9
Impulse Response Functions - France

U.S. 1961*- 2018

Var.	Shock	Horizon						
		1	2	4	12	24	48	60
π	ε_π	0,27	0,29	0,39	0,51	0,62	0,74	0,77
	ε_R	0,49	0,46	0,36	0,30	0,24	0,17	0,15
	ε_Y	0,24	0,25	0,25	0,19	0,14	0,09	0,08
R	ε_π	0,79	0,82	0,85	0,87	0,91	0,95	0,96
	ε_R	0,15	0,14	0,13	0,11	0,08	0,04	0,03
	ε_Y	0,06	0,04	0,02	0,02	0,01	0,01	0,01
Y	ε_π	0,14	0,17	0,19	0,05	0,02	0,01	0,01
	ε_R	0,10	0,07	0,04	0,01	0,02	0,01	0,01
	ε_Y	0,76	0,76	0,77	0,94	0,96	0,98	0,98

Japan 1961*- 2018

Var.	Shock	Horizon						
		1	2	4	12	24	48	60
π	ε_π	0,11	0,16	0,21	0,35	0,49	0,63	0,68
	ε_R	0,53	0,52	0,50	0,45	0,38	0,28	0,24
	ε_Y	0,35	0,33	0,29	0,19	0,13	0,09	0,08
R	ε_π	0,89	0,90	0,93	0,98	0,99	1,00	1,00
	ε_R	0,10	0,09	0,06	0,02	0,01	0,00	0,00
	ε_Y	0,01	0,01	0,01	0,00	0,00	0,00	0,00
Y	ε_π	0,02	0,03	0,04	0,02	0,01	0,01	0,01
	ε_R	0,29	0,32	0,30	0,17	0,09	0,05	0,04
	ε_Y	0,68	0,64	0,67	0,81	0,89	0,95	0,96

France 1961 -2018

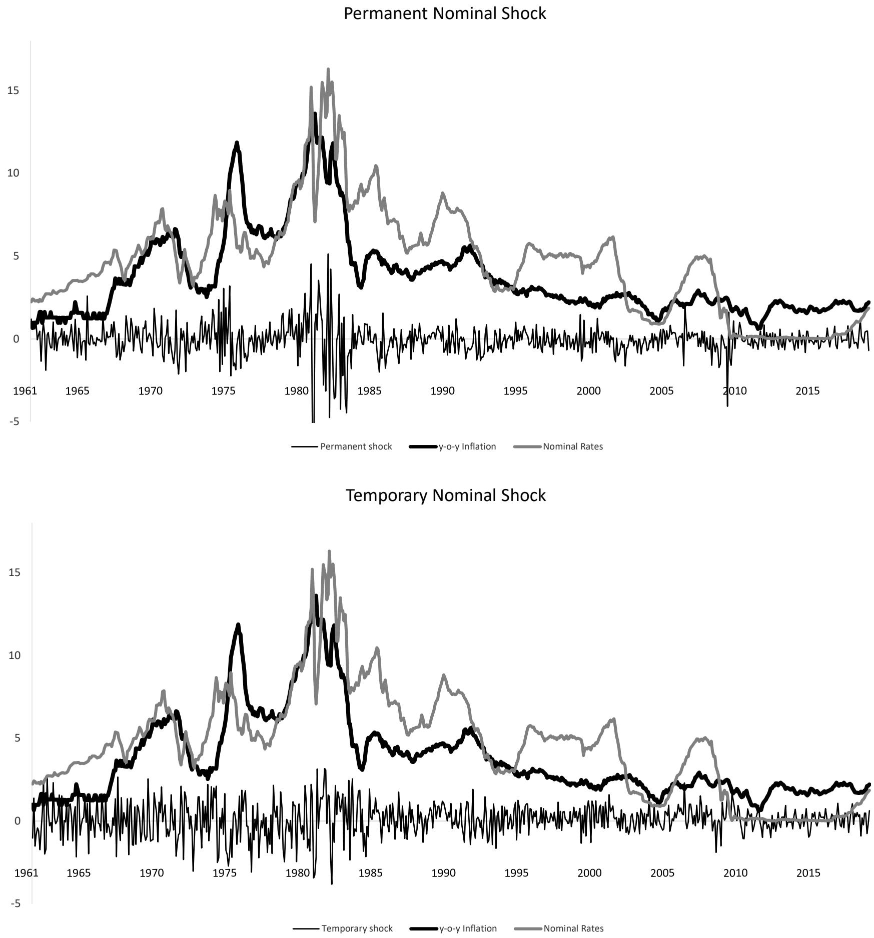
Var.	Shock	Horizon						
		1	2	4	12	24	48	60
π	ε_π	0,72	0,73	0,77	0,90	0,95	0,97	0,98
	ε_R	0,26	0,24	0,22	0,09	0,05	0,02	0,02
	ε_Y	0,03	0,03	0,02	0,01	0,01	0,00	0,00
R	ε_π	0,39	0,42	0,43	0,46	0,54	0,68	0,73
	ε_R	0,61	0,57	0,55	0,50	0,42	0,29	0,24
	ε_Y	0,00	0,01	0,02	0,04	0,04	0,03	0,00
Y	ε_π	0,01	0,01	0,01	0,01	0,00	0,00	0,00
	ε_R	0,01	0,01	0,01	0,01	0,01	0,01	0,01
	ε_Y	0,98	0,98	0,98	0,98	0,98	0,99	0,99

Forecast Error Variance decomposition from VECM model. Estimation samples for the U.S. is 1961 M1 - 2018 M5 (T=689), for Japan 1960 M8 - 2017 M6 (T=683) and for France 1961 M8 - 2018 M12 (T=689).

Table 9. Forecast Error Variance Decomposition

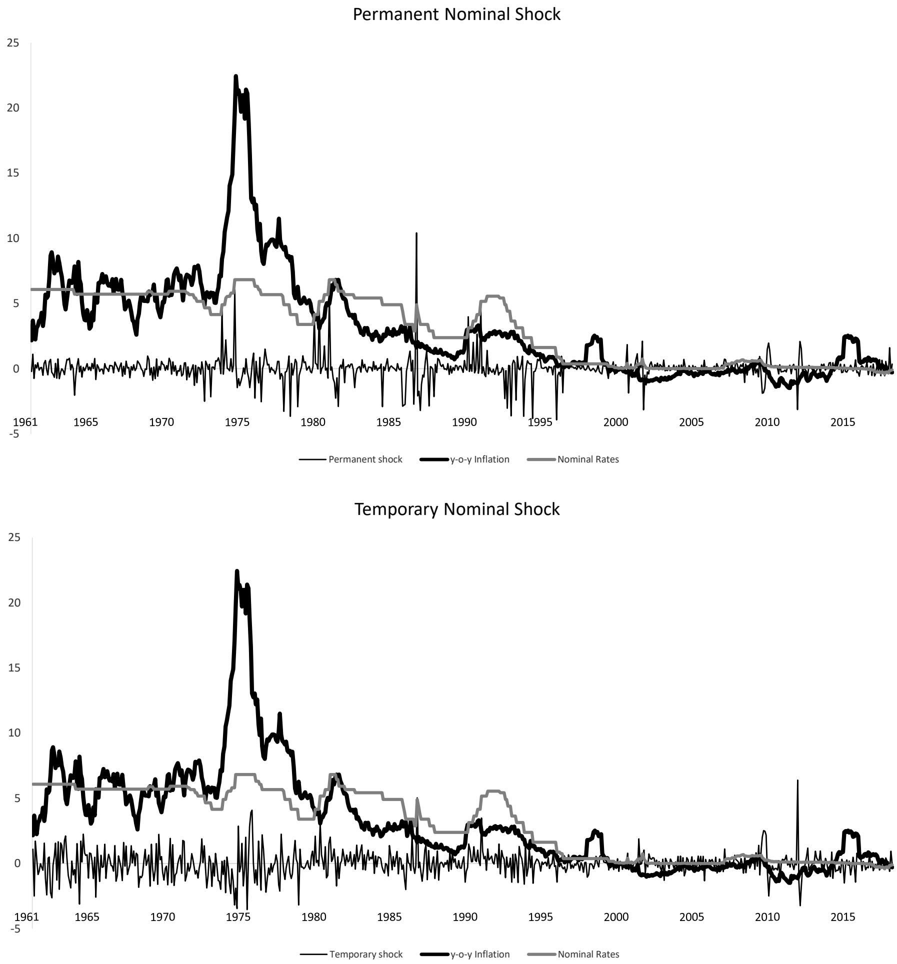
We now turn to the identified structural nominal shocks (Figures 10, 11 and 12): the permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Shocks are identified as described above and are such that $u_t = B\varepsilon_t$, where $\varepsilon_t = (\varepsilon_{\pi,t}, \varepsilon_{R,t}, \varepsilon_{Y,t})'$ is the vector of serially and mutually uncorrelated structural shocks with normalized (unit) variance and $u_t = (u_{\pi,t}, u_{R,t}, u_{Y,t})'$ is the vector of reduced form innovations. Given the normalization of the variance, the analysis of these plots ought to be complemented with the results for the forecast error variance decomposition. E.g., large identified temporary shocks may contribute little to the dynamics of nominal rates or inflation if their contribution to the variance of the forecast errors is small. For the three economies we report inflation and nominal interest rate data along with the shocks identified using the full sample. The appendix contains shocks identified using data for each of the sub-samples (pre-IT and post-IT). We highlight that the persistent shifts in inflation and nominal rates in the earlier part of the sample are clearly attributable to the identified permanent shocks (disinflation in the early to mid 1980's in the U.S. and France is attributed to large permanent and negative shocks), although relevant temporary shocks are also present (quite noticeably when spikes in nominal rates are quickly reversed). Permanent shocks also play a relevant role over the more recent period. Interestingly, the gradual shift associated with hitting the effective lower bound in the U.S., Japan and to a lesser extent in France , is attributable to permanent negative shocks (although large temporary shocks are also observed). In this regard results for Germany and euro area as a whole are more decisive if one is particularly interested in the euro area, in the sense that the recent period of low inflation and low nominal rates is clearly attributable to a large permanent shock.[24]

24. For the U.K., Germany and the euro area we find again clear similarities to what obtained for the U.S., France and Japan. For the U.K. and Germany, the persistent shifts in inflation and nominal rates in the earlier part of the sample are again clearly attributable to the identified permanent shocks. Also, the observed disinflation in the early to mid 1980's in the U.K. and Germany can be attributed to large permanent and negative shocks. Relevant temporary shocks are also present in the earlier part of the sample. We also notice that the volatility of the permanent shocks tends to be lower in the later part of the sample. The same is true for the temporary shocks but only in the case of the U.K. Permanent shocks play again a relevant role over the more recent period. The recent effective lower bound episodes in the U.K., Germany and also the euro area, are attributable to a large permanent negative shock (although relevant temporary shocks are also observed during this period).



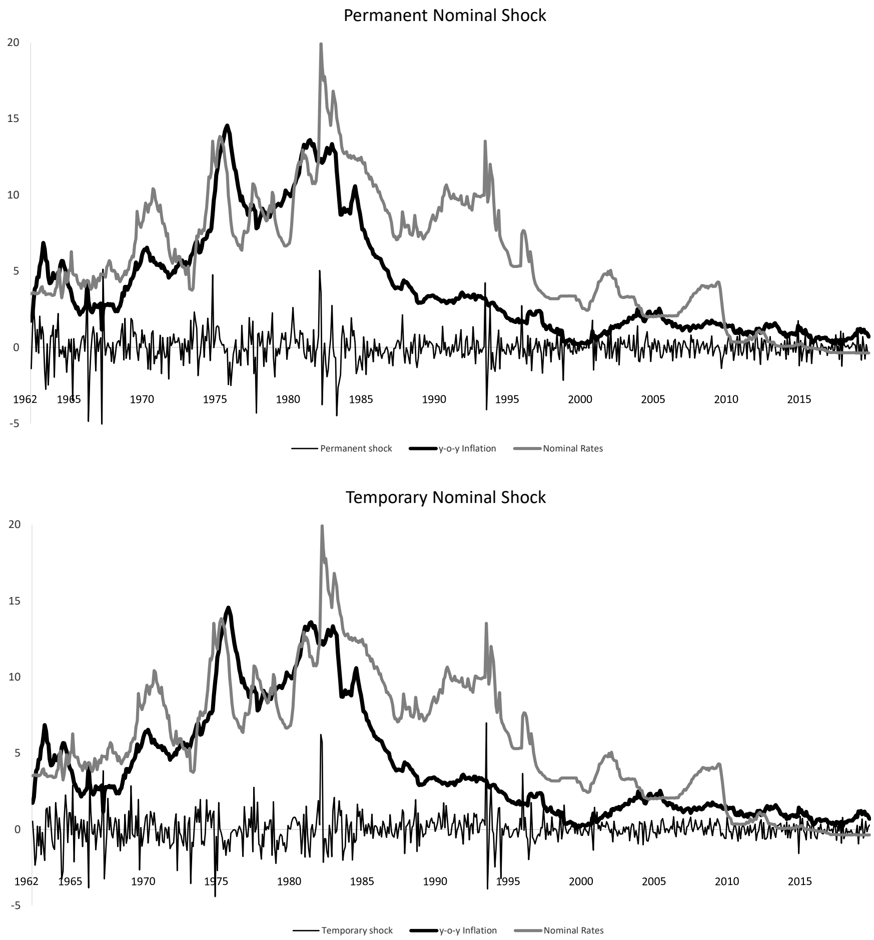
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 10
Identified Nominal Shocks - U.S., Full Sample



Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 11
Identified Nominal Shocks - Japan, Full Sample



Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 12
Identified Nominal Shocks - France, Full Sample

5. Conclusions

We show that a cross-section of averaged data is unambiguous about a positive, virtually one-for-one relationship between inflation and nominal interest rates. A simple estimation of this relationship yields a lower coefficient in the later part of the sample, one characterized by (successful) inflation targeting in most countries at low common values. Still, the fit to the line with unit slope is as good as the one obtained using the whole sample, i.e., the one-for-one relation can be a feature of the data, but harder to identify. Panel evidence suggests cointegration between the two variables, with a close to, but below, unitary coefficient under various specifications, even if one considers samples with marked shifts in inflation and nominal rates. A coefficient around 0.8 on nominal rates seems to better characterize the relation. This could be due to the slow fall in real interest rates observed over the last decades, which deserves further investigation. Within this panel setting we also report a substantially lower coefficient in the later part of the sample, which could be a result of a lack of clear shifts in both inflation and nominal rates over the more recent period. Our time series analysis conveys similar messages.

The difficulties in testing long-run relations, and monetary neutrality in particular, using reduced form statistical models are well-known at least since Lucas (1972) and Sargent (1973). Permanent shifts in the relevant time series, likewise a unit-root behavior, help overcome those difficulties and often dispense a detailed knowledge of the model generating the data, see Fisher and Seater (1993) or King and Watson (1997). It it thus no surprise that the results above can be reconciled with standard New Keynesian models where the Fisher relation is embedded, specially when models with stationary inflation and nominal rates are contrasted with versions displaying a unit-root in the two variables (and cointegration). This unit-root extension also allows the joint analysis of the effects of permanent and temporary monetary shocks.

We made an empirical attempt at distinguishing these shocks by estimating, for the U.S., Japan, France, the U.K., Germany and the euro area, a structural vector error correction model featuring cointegration between inflation and nominal rates, identified using standard long-run neutrality restrictions. We find that the effects of temporary nominal shocks are the standard ones, i.e., a temporary rise in nominal rates leads to a temporary fall in inflation and output; these effects dissipate in 3 years. More interesting is the effect of permanent nominal shocks. Given cointegration between inflation and nominal rates - whose coefficient we do not impose to be one -, permanent shocks lead to permanent shifts in nominal rates and inflation; after 3 years the adjustment is essentially made. More importantly, the short-run effects are similar to the long-run effects in that inflation responds positively and immediately to a permanent rise in nominal rates, i.e. we also find the "Neo-Fisher" effect first documented by Uribe (2017, 2018).

Overall, and qualitatively, the results are remarkably similar to what obtains in the aforementioned New Keynesian model with non-stationary inflation and nominal rates. This suggests that conventional effects and effects of permanent shocks can live together in models and in the data. Not acknowledging this distinction can lead to misleading identification of monetary shocks, specially given the fact that permanent shocks seem to account for a significant fraction of the forecast error variance in inflation and nominal rates, even at short horizons.

We highlighted from the onset that Japan has lived a period of very low inflation and essentially zero nominal rates for almost two decades while unemployment stands at 2.4% as of April 2019. The solutions envisaged to get inflation back on track (target) amount to provide more stimulus in the form of lower long rates along with promises to keep short rates low for long(er), see Bank of Japan (2016a, 2016b). Further, BoJ explicitly adopted an "inflation overshooting commitment": "the Bank commits itself to expanding the monetary base until the year-on-year rate of increase in the observed consumer price index (CPI) exceeds the price stability target of 2 percent and stays above the target in a stable manner.". Nothing happened to inflation since then.[25] What's more puzzling is that BoJ estimates the natural real rate to be close to 0% while pledging to continue this policy "until the year-on-year rate of increase in the observed consumer price index (CPI) exceeds the price stability target of 2 percent and stays above the target in a stable manner". Reaching an inflation rate of 2% seems quite a challenge given these premises. Explaining the zero inflation and zero nominal rates outcome is quite straightforward given BoJ's 0% estimate for the natural real rate...

Whether this sort of inconsistency plays a role in the euro area is still an open question, but a pressing one. With nominal rates pegged at the lower bound it is hard to see why inflation should converge towards 2%. Increasing nominal rates while inflation is low may at some point be the only feasible option to raise inflation back to target, both in the euro area and in Japan. Assessing whether this is optimal in any sense is beyond the scope of this paper. What our analysis highlights, for positive analyses, is the importance of distinguishing between temporary and permanent monetary policy actions. We confirm the important result in Uribe (2017,2018) that a permanent nominal shock associated with higher nominal rates raises inflation also in the short run.

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25. At the same time, there seems to be no contradiction between the objective for short rates, long rates, and asset purchases. Indeed, if the target for short rates is zero, one can expect that long rates converge to zero and remain there, regardless of asset purchases. In Japan, long rates are already close to zero while it is hard to argue that anyone needs to be further convinced that short rates will indeed remain low for a long time.

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Appendix

Country	Code
Australia	AU
Austria	AT
Belgium	BE
Canada	CA
Chile	CL
Czech Republic	CZ
Denmark	DK
Estonia	EE
Finland	FI
France	FR
Germany	DE
Greece	GR
Hungary	HU
Iceland	IS
Ireland	IE
Israel	IL
Italy	IT
Japan	JP
Korea	KO
Mexico	MX
Netherlands	NL
New Zealand	NZ
Norway	NO
Poland	PL
Portugal	PT
Slovak Republic	SK
Slovenia	SI
Spain	ES
Sweden	SE
Switzerland	CH
Turkey	TR
United Kingdom	UK
United States	US
Brazil	BR
China	CN
Colombia	CO
India	IN
Indonesia	ID
Latvia	LV
Russia	RU
South Africa	ZA

Table 10. Country Codes

Series	Frequency	Country	Source	Sample Period	Description and Notes
P					
Consumer Price Index, All Items, 2010=100	Annual	All but DK, IS, IE	OECD	1960-2016	Start year: CL - 1970; EE - 1998; IL - 1970; HU - 1980; MX - 1969; NL - 1961; PL - 1989; SK - 1991; SI - 1980; BR - 1980; CN - 1993; CO - 1970;ID - 1968; LV - 1991; RU - 1992
Consumer Price Index, All Items, 2010=100	Annual	DK, IS, IE	OECD and IMF IFS	1960-2016	IMF IFS up to 1966 for DK and up to 1975 for IS and IE
R					
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	AT	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1968 onwards
Overnight Interest rate, Monetary Policy rate and T-bill rate	Annual	CL	OECD and IMF IFS	1996-2016	Overnight rate for 1996 and 1997; Monetary policy rate for 2012-2015
Monetary Policy Interest Rate and T-bill rate	Annual	DK	OECD and IMF IFS	1960-2016	Monetary policy rate up to 1975; T-bill rate (IMF) for 1976-1986; T-bill rate (OECD) afterwards
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	FI	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1987 onwards
T-bill rate	Annual	GR	OECD and IMF IFS	1974-2016	T-bill rate (OECD) from 1995 onwards
T-bill rate	Annual	HU	OECD and IMF IFS	1991-2016	IMF for 2004-2007; 2009 and 2012-2013
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	IS	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1988 onwards
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	IE	OECD and IMF IFS	1960-2016	Discount rate for 1960-1969; IMF T-bill rate for 1970-1983; OECD T-bill rate afterwards
Nominal interest rate: T-bill rate	Annual	IL	OECD and IMF IFS	1984-2016	T-bill rate (OECD) from 1992 onwards
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	IT	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1979 onwards
Nominal Interest Rate: T-bill rate	Annual	JP	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 2003 onwards
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	KO	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1991 onwards
Nominal Interest Rate: Overnight Interest rate and T-bill rate	Annual	MX	OECD and IMF IFS	1978-2016	IMF T-bill rate for 1978-1985 and 1987-1996; OECD overnight rate for 1986; T-bill rate (OECD) from 1997 onwards
Monetary Policy rate and T-bill rate	Annual	NL	OECD and IMF IFS	1960-2016	Monetary policy rate up to 1975; T-bill rate (IMF) for 1976-1985; T-bill rate (OECD) afterwards
Monetary Policy rate and T-bill rate	Annual	NO	OECD and IMF IFS	1960-2016	Monetary policy rate up to 1978; T-bill rate (OECD) afterwards
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	PT	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1999 onwards
Nominal Interest Rate: T-bill rate	Annual	SI	OECD and IMF IFS	1999-2016	T-bill rate (OECD) from 2002 onwards
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	ES	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1977 onwards
Nominal Interest Rate: T-bill rate	Annual	SE	OECD and IMF IFS	1963-2016	T-bill rate (OECD) from 1981 onwards
Monetary Policy rate and T-bill rate	Annual	CH	OECD and IMF IFS	1960-2016	Monetary policy rate up to 1973; T-bill rate (OECD) afterwards
Nominal Interest Rate: Overnight Interest rate	Annual	TR	OECD	1986-2016	
Nominal Interest Rate: T-bill rate	Annual	UK	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1978 onwards
Nominal Interest Rate: T-bill rate	Annual	US	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1965 onwards
Nominal Interest Rate: T-bill rate	Annual	BR	IMF IFS	1995-2016	
Nominal Interest Rate: Overnight Interest rate	Annual	CN	OECD	1990-2016	
Nominal Interest Rate: Discount Rate and T-bill rate	Annual	CO	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1986 onwards
Nominal Interest Rate: Discount Rate	Annual	IN	IMF IFS	1963-2016	
Nominal Interest Rate: T-bill rate	Annual	ZA	OECD and IMF IFS	1960-2016	T-bill rate (OECD) from 1981 onwards
Nominal Interest Rate: T-bill rate	Annual	All others	OECD	1960-2016	Start year: AU - 1968; CZ - 1993; EE - 1996; FR - 1970; NZ - 1974; PL - 1992; SK - 1999; ID - 1998; LV - 1998; RU - 1997

Table 11. Data sources for annual data on inflation and short rates

Country	Year of IT adoption	Source and/or justification
AU	1993	Gill Hammond (2012)
AT	1999	Adoption of the Euro; no previous statement or low π
BE	1988	Stabilization of inflation at lower levels
CA	1991	Gill Hammond (2012)
CL	2000	Gill Hammond (2012)
CZ	2000	Stabilization of inflation at lower levels
DK	1995	Nominal interest rate reaches average of subsequent period
EE	2002	Preparation to join the Euro
FI	1993	Bank of Finland statement
FR	1999	Adoption of the Euro; no previous statement or low π
DE	1961	Very low inflation throughout the sample
GR	1999	Stabilization of inflation at lower levels
HU	2001	Gill Hammond (2012)
IS	2001	Gill Hammond (2012)
IE	1999	Adoption of the Euro; no previous statement or low π
IL	1997	Gill Hammond (2012)
IT	1999	Adoption of the Euro; no previous statement or low π
JP	2001	Gill Hammond (2012)
KO	1998	Gill Hammond (2012)
MX	2001	Stabilization of inflation at lower levels
NL	1999	Gill Hammond (2012)
NZ	1989	Gill Hammond (2012)
NO	2001	Gill Hammond (2012)
PL	2001	Stabilization of inflation at lower levels
PT	1999	Adoption of the Euro; no previous statement or low π
SK	2002	Preparation to join the Euro
SI	2002	Preparation to join the Euro
ES	1994	Bank of Spain statement
SE	1993	Bank of Sweden statement
CH	1961	Very low inflation throughout the sample
TR	2006	Gill Hammond (2012)
UK	1992	Gill Hammond (2012)
US	1984	Onset of Great Moderation; common break point
BR	1998	Stabilization of inflation at lower levels
CN	1998	Stabilization of inflation at lower levels
CO	2000	Gill Hammond (2012)
IN	2000	Stabilization of inflation at lower levels
ID	2005	Gill Hammond (2012)
LV	2000	Stabilization of inflation at lower levels
RU	2004	Stabilization of inflation at lower levels
ZA	2000	Stabilization of inflation at lower levels

Table 12. Beginning of Inflation Targeting Dates

Description	FRED code
Consumer Price Index for All Urban Consumers: All Items Less Food and Energy	CPILFESL
Consumer Price Index: OECD Groups: All Items Non-Food and Non-Energy (Japan)	CPGRLE01JPM659N
Harmonized Index of Consumer Prices: Overall Index Excluding Energy, Food, Alcohol, and Tobacco for Euro area (EA11-2000, EA12-2006, EA13-2007, EA15-2008, EA16-2010, EA17-2013, EA18-2014, EA19)	00XEFDEZCCM086NEST
Consumer Price Index of All Items in France (only up to 1970 M12)	FRACPIALLMINMEI
Consumer Price Index: All Items Excluding Food and Energy for France (from 1971 M1 onwards)	FRACPICORMINMEI
Consumer Price Index of All Items in the United Kingdom (only up to 1970 M12)	GBRCPIALLMINMEI
Consumer Price Index: All Items Excluding Food and Energy for U.K. (from 1971 M1 onwards)	GBRCPICORMINMEI
Consumer Price Index: All Items Excluding Food and Energy for Germany	DEUCPICORMINMEI
Industrial Production Index (U.S.)	INDPRO
Production of Total Industry in Japan (Industrial Production Index)	JPNPROINDMISMEI
Euro area 19 (fixed composition) - Industrial Production Index, Total Industry - NACE Rev2	[Statistical Data Warehouse -ECB]
Production of Total Industry in France	FRAPROINDMISMEI
Production of Total Industry in the United Kingdom	GBRPROINDMISMEI
Production of Total Industry in Germany	DEUPROINDMISMEI
3-Month Treasury Bill: Secondary Market Rate (U.S.)	TB3MS
Interest Rates, Government Securities, Treasury Bills for Japan	INTGSTJPM193N
3-Month or 90-day Rates and Yields: Interbank Rates for the Euro Area	IR3TIB01EZM156N
Immediate Rates: Less than 24 Hours: Call Money/Interbank Rate for France	IRSTCI01FRM156N
3-Month or 90-day Rates and Yields: Treasury Securities for the U.K.	IR3TTS01GBM156N
3-Month or 90-day Rates and Yields: Interbank Rates for Germany	IR3TIB01DEM156N

Note: sources are Federal Reserve Bank of St. Louis's FRED and ECB's Statistical Data Warehouse

Table 13. Monthly data for core inflation, short rates and industrial production

Model with varying inflation target

Households

Households' problem is to choose sequences of consumption, C_t, and labour, L_t, that maximize:

$$E_0 \sum_{t=0}^{\infty} A_t \beta^t U(C_t, L_t) \tag{14}$$

where A_t is a preference shock, $U(C_t, L_t) = ln(C_t) - \frac{L_t^{1+\varphi}}{1+\varphi}$ and C_t is a Dixit-Stiglitz aggregator of consumption varieties $j \in [0,1]$, i.e., $C_t = \left(\int_0^1 C_t(j)^{1-\frac{1}{\varepsilon}} dj\right)^{\frac{\varepsilon}{\varepsilon-1}}$ where ε is the elasticity of substitution across varieties. The budget constraint of households reads as:

$$P_t C_t + B_t = R_{t-1} B_{t-1} + W_t L_t \tag{15}$$

where R_t is the gross nominal interest rate on one period bonds, B_t, and $P_t = \left[\int_0^1 P_t(j)^{1-\varepsilon} dj\right]^{\frac{1}{1-\varepsilon}}$ is the aggregate price level given the choices of households over varieties, where $P_t(j)$ is the price of one unit of variety j and W_t is the wage rate. Intratemporal optimization (choice between labor and consumption) yields:

$$\frac{W_t}{P_t} = L_t^{\varphi} C_t \tag{16}$$

Intertemporal optimization (choice between consumption and saving) yields the Euler Equation:

$$A_t C_t^{-1} = E_t(A_{t+1}) \beta (1 + R_t) E_t(C_{t+1}^{-1} \Pi_{t+1}^{-1}) \tag{17}$$

where Π_{t+1} is the gross inflation rate $\frac{P_{t+1}}{P_t}$.

Firms

The only difference in our firm problem relative to Juillard et al. (2008) is that we do not have capital in our model. Our expression for marginal cost is different, but all the rest will be similar.A firm that is allowed to reoptimize its price (this occurs with probability $1-\delta$) chooses a current price V_t and a growth rate of price υ_t at which to update the price from today until the time it is able to change its policy. If at time $t+k$ the firm keeps its time t price (i.e., it could not reoptimize from t to $t+k$), its price is therefore $P_{t+k} = V_t {\upsilon_t}^k$. A generic firm maximizes:

$$E_t \sum_{k=0}^{\infty} (\delta\beta)^k \lambda_{t+k} \left[\Big(\frac{V_t \upsilon_t^k}{P_{t+k}}\Big)^{1-\varepsilon} Y_{t+k} - MC_{t+k} \Big(\frac{V_t (\upsilon_t)^k}{P_{t+k}}\Big)^{-\varepsilon} \right] \tag{18}$$

subject to demand by households. λ_{t+k} is the marginal utility of consumption, Y_{t+k} is output and MC_t is the marginal cost, given by:

$$MC_t = \frac{W_t}{1-\alpha} Y_t^{\frac{\alpha}{1-\alpha}} \tag{19}$$

taking into account the production function $Y_t = L_t^{1-\alpha}$. This delivers prices as a markup μ_t over marginal cost.

Linearized equations of the model

We adapt the equations in Juillard et al. (2008) as needed. All variables are in log deviations from the steady-state (denoted with lower cases and hat). The nominal interest rate (R_t) and the inflation rate(π_t) are in log deviations from the target inflation rate (denoted $\widehat{R}_t$ and $\hat{\pi}_t$). $\triangle E_t(\hat{a}_{t+1})$ is denoted by ε_t^{β}, a preference shock. The target rate is denoted by $\overline{\pi}_t$. The shocks to the model are ε_t^m, $\varepsilon_t^{\overline{\pi}}$, ε_t^{β} and ε_t^c (put directly in the Phillips curve), assumed i.i.d. in our experiments (but they could be serially correlated).
Euler equation:

$$\hat{y}_t = E_t(\hat{y}_{t+1}) - (\widehat{R}_t - E_t(\hat{\pi}_{t+1})) - \varepsilon_t^{\beta} \tag{20}$$

Phillips curve with auxiliary variables ψ_t,υ_t:

$$\begin{aligned} E_t(\hat{\pi}_{t+1}) = \hat{\pi}_t\Big(\frac{2}{\beta} - \delta\Big) + \hat{\upsilon}_t((1-\delta)(1+\delta)) + \hat{\psi}_t\Big(\delta(1+\delta) - \frac{2}{\beta}\Big) + \varepsilon_t^c \\ - \frac{2(1-\delta)(1-\delta\beta)}{\delta\beta}(\hat{mc}_t + \hat{\mu}_t) + (1-\delta)(E_t(\hat{\mu}_{t+1}) - \hat{\mu}_t) \end{aligned} \tag{21}$$

$$E_t(\hat{\upsilon}_{t+1}) = \hat{\upsilon}_t + \frac{(1-\delta\beta)^2}{(\delta\beta)^2}\frac{\delta}{1-\delta}\hat{\psi}_t - \frac{(1-\delta\beta)^2}{(\delta\beta)^2}\frac{\delta}{1-\delta}\hat{\pi}_t + \frac{(1-\delta\beta)^2}{(\delta\beta)^2}(\hat{mc}_t + \hat{\mu}_t) \tag{22}$$

$$\hat{\psi}_t = \delta\hat{\psi}_{t-1} + (1-\delta)\hat{\upsilon}_{t-1} - \overline{\pi}_t \tag{23}$$

Marginal cost:

$$\hat{mc}_t = \frac{1+\varphi}{1-\alpha}\hat{y}_t \tag{24}$$

Inflation target:

$$\overline{\pi}_t = \overline{\pi}_{t-1} + \varepsilon_t^{\overline{\pi}}$$

Taylor rule (version without output under flexible prices):

$$\widehat{R}_t = \rho(\widehat{R}_{t-1} - \overline{\pi}_t) + (1-\rho)(\varphi_\pi\hat{\pi}_t + \varphi_y\hat{y}_t + \varphi_{\triangle y}(\hat{y}_t - \hat{y}_{t-1})) + \varepsilon_t^m \tag{25}$$

The Neutrality of Nominal Rates: How Long is the Long Run?
APPENDIX - FURTHER RESULTS

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Abstract
This appendix contains more complete results regarding the estimation of the structural vector error correction model (SVECM). For the U.S., Japan and France we report impulse response functions for the pre-inflation targeting period, the forecast error variance decomposition for sub-samples and the identified structural nominal shocks (the temporary and the permanent). We repeat the whole analysis for the U.K., Germany and the euro area.

Note: The views here are entirely our own and not necessarily those of Banco de Portugal or the Eurosystem. Corresponding author Pedro Teles: E-mail: pteles@ucp.pt Address: Av. Almirante Reis 71 6th floor 1150-012 Lisboa, Portugal.

1. Structural VECM - Review

We aim at identifying permanent and transitory nominal shocks. Permanent shocks ought to drive the long-run positive relation between inflation and nominal rates, since the two variables can be described as cointegrated whenever they are better characterized by a unit-root. As in the main text, for the estimation of the SVECM we will assume inflation and nominal rates are cointegrated (the coefficient does not need to be one) and we impose that output, as measured by industrial production, is not cointegrated with those two variables. Assuming a vector autoregression representation for the data, if the three variables are integrated of order one and cointegrated they can be represented in an error correction framework. The VECM reduced form representation is given by:

$$\Delta X_t = \alpha_0 + +\gamma\beta X_{t-1} + \sum_{j=1}^{K_r} \beta_j \Delta X_{i,t-j} + u_t$$

$X_t = (\pi_t, R_t, Y_t)'$ collects inflation, nominal rates and output and $\gamma := (\gamma_\pi, \gamma_R, \gamma_Y)'$ collects the adjustment coefficients. $\beta := (1 : -\delta : 0)$ is the cointegrating vector, where we make explicit that output is not part of the cointegration relation. $u_t = (u_{\pi,t}, u_{R,t}, u_{Y,t})'$ is a vector of reduced form innovations. As in Lütkepohl (2006, Chapter 9), for identification of structural shocks we assume these are related to reduced form shocks through some non-singular matrix B such that $u_t = B\varepsilon_t$, where $\varepsilon_t = (\varepsilon_{\pi,t}, \varepsilon_{R,t}, \varepsilon_{Y,t})'$ is a vector of serially and mutually uncorrelated structural shocks. Restrictions on B are imposed to identify the structural shocks and assess their effects. The reduced form implies a Beveridge-Nelson decomposition which reads in structural form as:

$$X_t = X_{0,t}^* + \Xi B \sum_{i=1}^{t} \varepsilon_i + \sum_{j=0}^{\infty} \Xi_j^* B \varepsilon_{t-j} \tag{1}$$

where the last term is absolutely summable and thus stationary. $X_{0,t}^*$ contains the initial values and possibly linear trends while the second term is the so-called stochastic trend. Matrix Ξ is a straightforward function of the parameters of the reduced form and it has reduced rank if there is cointegration. We are assuming a cointegration rank of one so Ξ has rank 2. Since B is a non-singular matrix then ΞB will have rank 2 as all the variables in X_t are integrated but π_t and R_t are cointegrated, thus sharing a stochastic trend. As a first step to identify B one can set one of the columns of ΞB to zero. This should not be the third column (the loadings related to $\sum_{i=1}^{t} \varepsilon_{Y,i}$) since output should retain a stochastic trend, which is not, by assumption, the stochastic trend shared by π_t and R_t. Setting the first or second column of ΞB to zero amounts to establishing which of the shocks $\varepsilon_{\pi,t}$ or $\varepsilon_{R,t}$ is the temporary (as opposed to the permanent) nominal shock. We will always pick $\varepsilon_{\pi,t}$ as the permanent shock and $\varepsilon_{R,t}$ as the transitory shock; the second column of ΞB is thus set to zero. Thus, only $\varepsilon_{\pi,t}$ will have a permanent effect on π_t and R_t.[1] We need one extra restriction to identify the permanent shock to output, or locally identify B. It suffices to place restrictions on ΞB. We assume that the permanent output shock does not affect the long-run level of both inflation and/or nominal rates (the first two elements in the third column of ΞB are set to zero, but restricting only one of these to zero would result in an equivalent identification). Finally, we overidentify the model and assume that the permanent nominal shock does not affect output in the long run.[2]

2. Long-run and short-run dynamics in the U.S., Japan and France

2.1. Pre inflation targeting IRFs

We report here the impulse response functions of the identified structural shocks for the U.S., Japan and France for the pre-inflation targeting period (Figures 1, 2 and 3). We focus on nominal shocks. As in the full sample, the effects of the temporary nominal shock are the conventional ones: a temporary shock increases nominal rates and drives inflation downwards. In the case of Japan the moderate increase in nominal rates is associated with a strong fall in inflation. Output falls on impact in the U.S. and Japan. In the case of Japan the fall in output is more persistent. Turning now to the permanent nominal shock (leading to a permanent increase in nominal rates) we notice that, as expected, the permanent increase in nominal rates is accompanied by an increase in inflation towards the new long-run level. Again, this adjustment takes place in roughly less than two years in all cases while qualitatively the short-run effect of this permanent shock is very similar to the long-run one. No fall of inflation in the short run is observed. In a nutshell, the impulse response functions

For a post-IT sample the impulse response functions for the U.S. and France would be similar but imprecisely estimated (wide confidence bands), since the shifts in inflation and nominal rates are modest. For Japan the impulse response functions are quite meaningless since cointegration does not obtain in the post-IT period (as inflation and nominal rates are well characterized by stationarity, recall also the estimation results in the main text).

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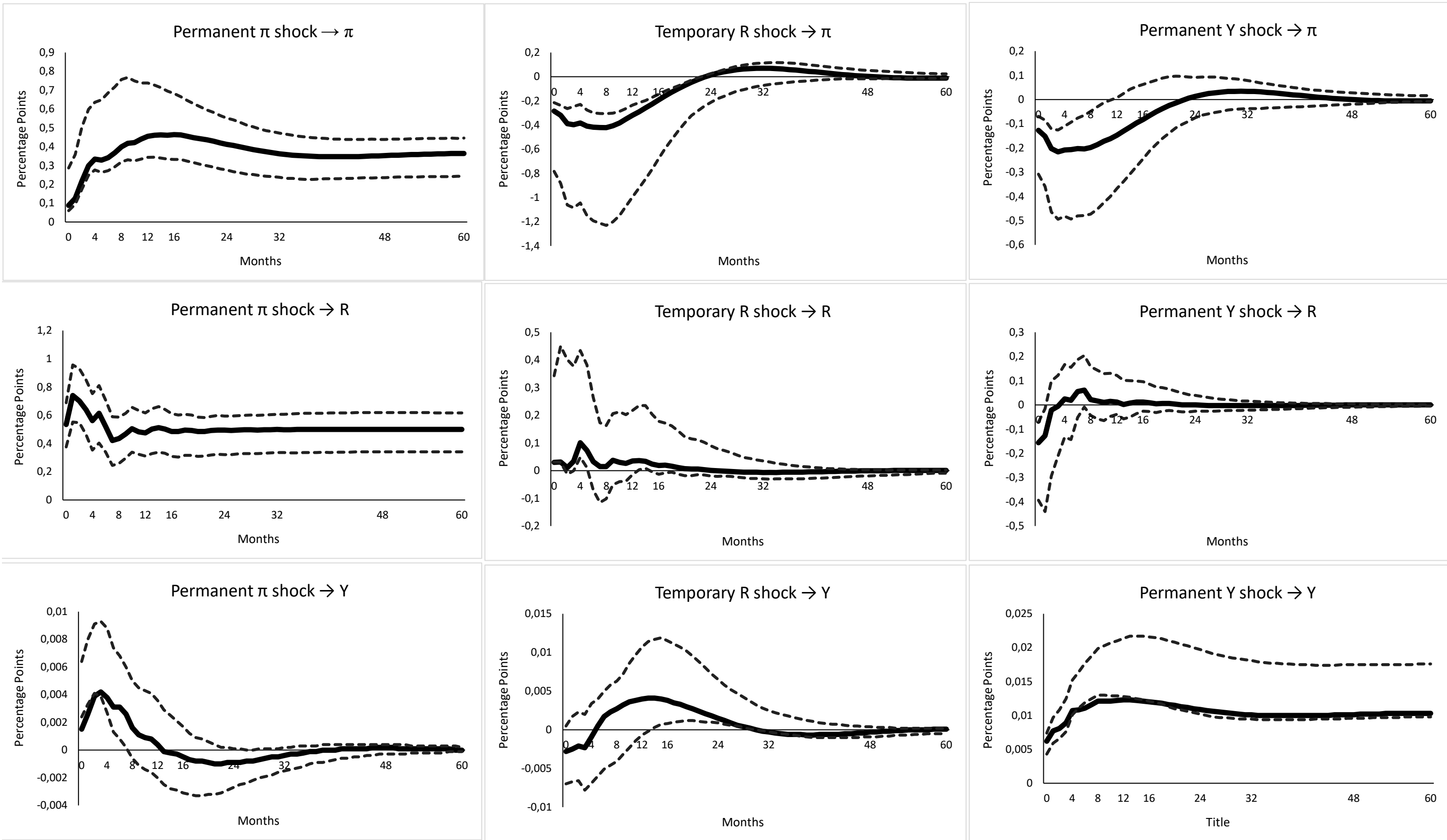
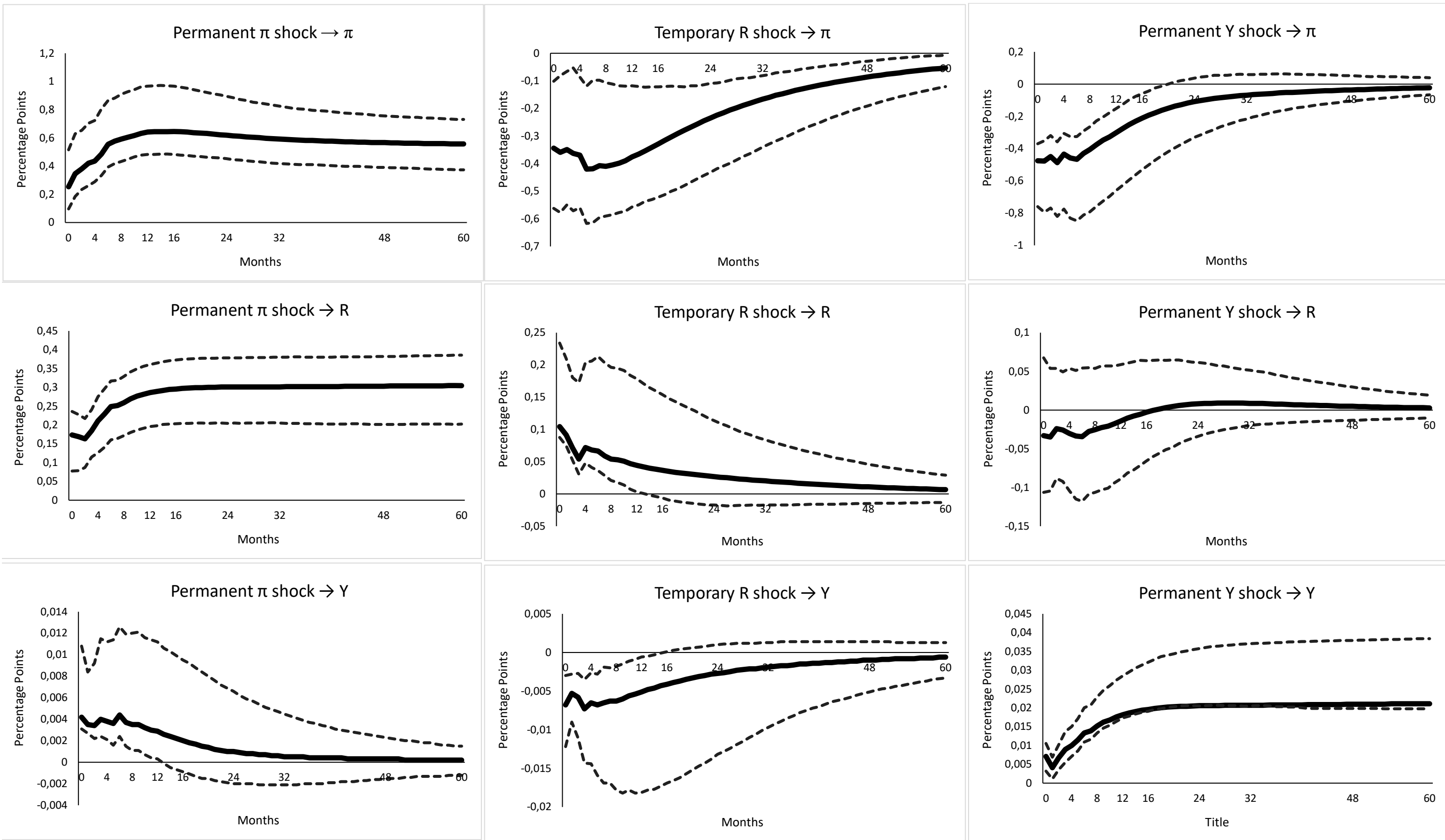
1. Given that the permanent shock only has permanent effects on inflation and nominal rates, this choice is innocuous. We could alternatively pick $\varepsilon_{R,t}$ as the permanent shock and $\varepsilon_{\pi,t}$ as the transitory shock; the first column of ΞB would be set to zero such that only $\varepsilon_{R,t}$ would have a permanent effect on π_t and R_t. In this way we would obtain exactly the same impulse response functions as those obtained with the alternative identification, only the labeling of the shocks would be switched ($\varepsilon_{R,t}$ would be the permanent shock).

2. Empirically, not adding this extra assumption changes little the results since the estimated entry in ΞB is very close to zero. Alternatively, one could only set this restriction and allow for permanent effects of output shocks on inflation and nominal rates. We should notice that this would only affect somewhat the impulse response functions associated with the output shock, not the ones associated with the nominal shocks. Results are available upon request.



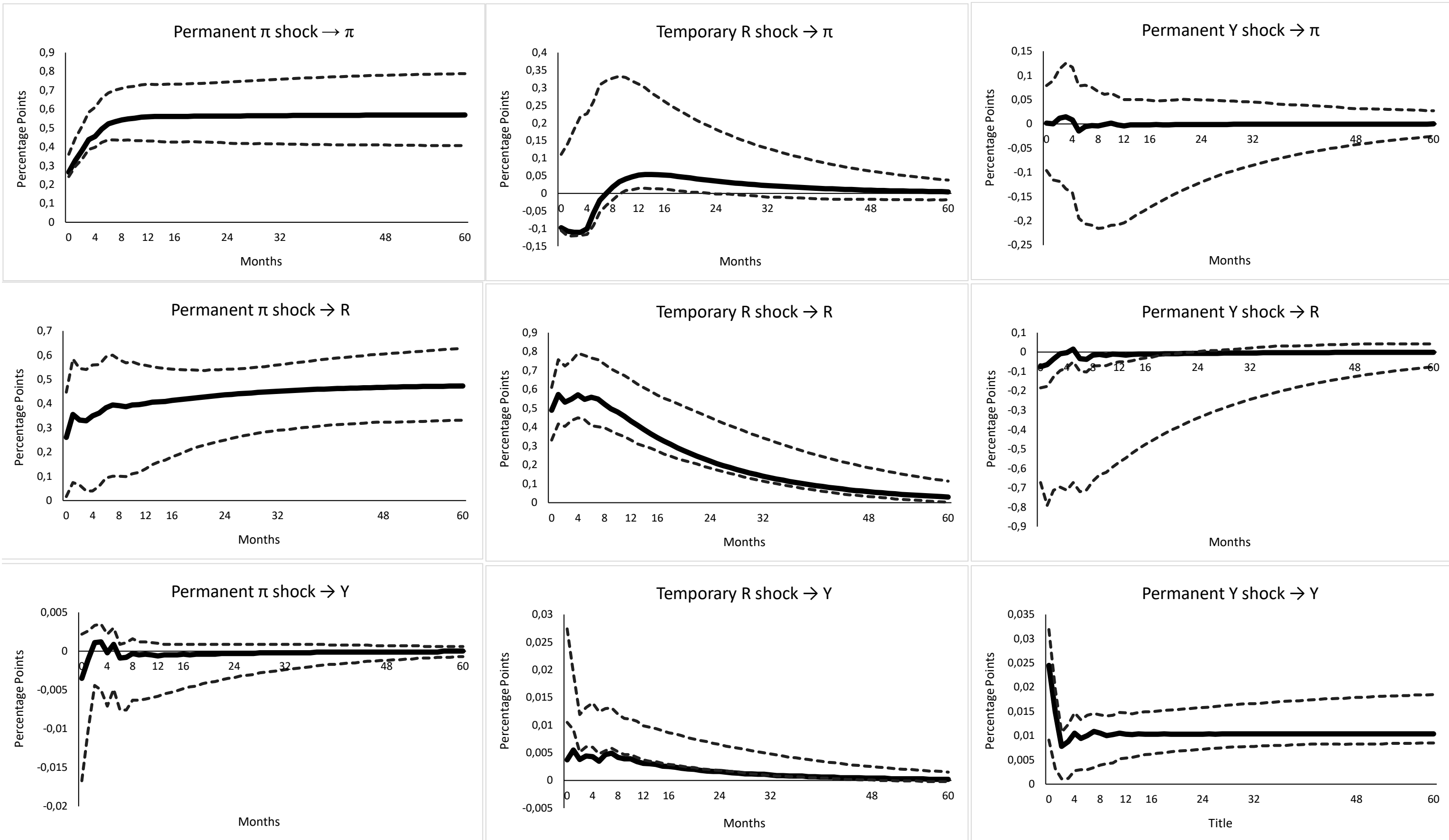
Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 1
Impulse Response Functions - U.S., Pre-IT



Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 2
Impulse Response Functions - Japan, Pre-IT



Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 3
Impulse Response Functions - France, Pre-IT

2.2. Forecast Error Variance Decomposition - Full Results

Next we analyze the decomposition of the variance of the forecast error for π_t, R_t and Y_t. Tables 1,2 and 3 display the results for the various sub-samples. Several patterns are stable across countries and samples. First, the bulk of the variance of the forecast error in the case of output is attributed to the output shock, even at short horizons. This shock also accounts for a significant fraction of the variance of the forecast error of inflation in several instances. In the case of Japan this is not true for the post-IT subsample whereas for France it is only true over the more recent sub-sample. Next, the contribution of the permanent nominal shock for the variance of the forecast error of nominal rates is always quite high, even at short horizons (less so in the case of France in the pre-IT sample). This shock also accounts for an important fraction of the variance of the forecast error for inflation, and very clearly so at horizons greater than two years. Interestingly, this contribution decreases somewhat at short horizons when the focus in on the post-IT sample. As for the temporary nominal shock, it often contributes relevantly to the variance of the forecast error of inflation at short horizons, and more so in the post-IT sample for France and Japan. It also contributes somewhat to the variance of the forecast error of nominal rates (most notably for France and less importantly in the U.S.). The contribution of this temporary shock to the variance of the forecast error of output is generally rather low, except for Japan in the pre-IT and full samples.

US 1961-2018

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,27	0,29	0,39	0,51	0,62	0,74	0,77
	ε_R	0,49	0,46	0,36	0,30	0,24	0,17	0,15
	ε_Y	0,24	0,25	0,25	0,19	0,14	0,09	0,08
R	ε_π	0,79	0,82	0,85	0,87	0,91	0,95	0,96
	ε_R	0,15	0,14	0,13	0,11	0,08	0,04	0,03
	ε_Y	0,06	0,04	0,02	0,02	0,02	0,01	0,00
Y	ε_π	0,14	0,17	0,19	0,05	0,02	0,01	0,01
	ε_R	0,10	0,07	0,04	0,01	0,02	0,01	0,01
	ε_Y	0,76	0,76	0,77	0,94	0,96	0,98	0,98

US 1961-1984

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,07	0,10	0,20	0,38	0,59	0,72	0,76
	ε_R	0,77	0,75	0,63	0,50	0,34	0,23	0,20
	ε_Y	0,15	0,16	0,16	0,12	0,08	0,05	0,05
R	ε_π	0,92	0,95	0,98	0,98	0,99	0,99	0,99
	ε_R	0,00	0,00	0,00	0,01	0,00	0,00	0,00
	ε_Y	0,08	0,05	0,02	0,01	0,01	0,00	0,00
Y	ε_π	0,05	0,08	0,14	0,06	0,03	0,02	0,01
	ε_R	0,16	0,11	0,08	0,05	0,06	0,04	0,03
	ε_Y	0,79	0,81	0,79	0,89	0,91	0,95	0,95

US 1984-2018

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,20	0,24	0,30	0,39	0,51	0,67	0,73
	ε_R	0,21	0,16	0,14	0,10	0,06	0,04	0,03
	ε_Y	0,60	0,60	0,57	0,51	0,43	0,29	0,24
R	ε_π	0,64	0,65	0,67	0,70	0,74	0,83	0,86
	ε_R	0,01	0,01	0,02	0,03	0,02	0,01	0,01
	ε_Y	0,36	0,34	0,31	0,28	0,23	0,16	0,00
Y	ε_π	0,39	0,38	0,35	0,30	0,23	0,13	0,10
	ε_R	0,03	0,04	0,04	0,04	0,03	0,02	0,02
	ε_Y	0,58	0,58	0,61	0,66	0,74	0,85	0,88

Forecast Error Variance decomposition in VECM model. Estimation samples for the U.S: i) 1961-2018, or 1961 M1 - 2018 M5 (T=689), ii) 1961-1984, or 1961 M1 - 1984M1 (T=282) and iii) 1984-2018, or 1984 M1 -2018 M5 (T=413)

Table 1. Forecast Error Variance Decomposition - U.S.

Japan 1960-2017

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,11	0,16	0,21	0,35	0,49	0,63	0,68
	ε_R	0,53	0,52	0,50	0,45	0,38	0,28	0,24
	ε_Y	0,35	0,33	0,29	0,19	0,13	0,09	0,08
R	ε_π	0,89	0,90	0,93	0,98	0,99	1,00	1,00
	ε_R	0,10	0,09	0,06	0,02	0,01	0,00	0,00
	ε_Y	0,01	0,01	0,01	0,00	0,00	0,00	0,00
Y	ε_π	0,02	0,03	0,04	0,02	0,01	0,01	0,01
	ε_R	0,29	0,32	0,30	0,17	0,09	0,05	0,04
	ε_Y	0,68	0,64	0,67	0,81	0,89	0,95	0,96

Japan 1960-2001

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,16	0,21	0,27	0,43	0,58	0,72	0,76
	ε_R	0,29	0,28	0,26	0,25	0,22	0,16	0,14
	ε_Y	0,55	0,51	0,47	0,31	0,20	0,12	0,11
R	ε_π	0,72	0,73	0,79	0,91	0,95	0,98	0,98
	ε_R	0,26	0,24	0,18	0,08	0,04	0,02	0,02
	ε_Y	0,03	0,03	0,02	0,01	0,01	0,00	0,00
Y	ε_π	0,16	0,18	0,14	0,07	0,03	0,01	0,01
	ε_R	0,41	0,43	0,39	0,19	0,09	0,04	0,03
	ε_Y	0,44	0,39	0,47	0,75	0,88	0,95	0,96

Japan 2001-2017

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,19	0,23	0,25	0,22	0,24	0,33	0,37
	ε_R	0,81	0,77	0,74	0,78	0,75	0,66	0,62
	ε_Y	0,00	0,00	0,00	0,00	0,01	0,01	0,00
R	ε_π	0,67	0,69	0,74	0,87	0,93	0,96	0,97
	ε_R	0,31	0,30	0,26	0,13	0,07	0,03	0,03
	ε_Y	0,01	0,01	0,01	0,00	0,00	0,00	0,00
Y	ε_π	0,16	0,12	0,06	0,02	0,01	0,00	0,00
	ε_R	0,00	0,02	0,04	0,02	0,01	0,01	0,00
	ε_Y	0,84	0,86	0,89	0,95	0,98	0,99	0,99

Forecast Error Variance decomposition in VECM model. Estimation samples for Japan: i) 1961-2017, or 1960 M8 - 2017 M6 (T=683), ii) 1961-2001, or 1960 M8 - 2001 M12 (T=497) and iii) 2001-2017, or 2001 M1 -2017 M6 (T=198)

Table 2. Forecast Error Variance Decomposition - Japan

France 1961-2018

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,72	0,73	0,77	0,90	0,95	0,97	0,98
	ε_R	0,26	0,24	0,22	0,09	0,05	0,02	0,02
	ε_Y	0,03	0,03	0,02	0,01	0,01	0,00	0,00
R	ε_π	0,39	0,42	0,43	0,46	0,54	0,68	0,73
	ε_R	0,61	0,57	0,55	0,50	0,42	0,29	0,24
	ε_Y	0,00	0,01	0,02	0,04	0,04	0,03	0,00
Y	ε_π	0,01	0,01	0,01	0,01	0,00	0,00	0,00
	ε_R	0,01	0,01	0,01	0,01	0,01	0,01	0,01
	ε_Y	0,98	0,98	0,98	0,98	0,98	0,99	0,99

France 1961-1999

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,88	0,89	0,92	0,98	0,99	0,99	0,99
	ε_R	0,12	0,11	0,08	0,02	0,01	0,01	0,01
	ε_Y	0,00	0,00	0,00	0,00	0,00	0,00	0,00
R	ε_π	0,22	0,25	0,26	0,32	0,44	0,63	0,69
	ε_R	0,77	0,74	0,73	0,68	0,56	0,37	0,31
	ε_Y	0,02	0,01	0,01	0,00	0,00	0,00	0,00
Y	ε_π	0,02	0,02	0,02	0,01	0,01	0,00	0,00
	ε_R	0,02	0,05	0,07	0,10	0,08	0,05	0,04
	ε_Y	0,96	0,94	0,91	0,89	0,91	0,95	0,96

France 1999-2018

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,33	0,35	0,34	0,35	0,32	0,54	0,66
	ε_R	0,45	0,45	0,46	0,44	0,47	0,32	0,24
	ε_Y	0,21	0,20	0,20	0,21	0,20	0,13	0,10
R	ε_π	0,78	0,82	0,85	0,93	0,96	0,98	0,99
	ε_R	0,05	0,04	0,04	0,04	0,02	0,01	0,01
	ε_Y	0,18	0,14	0,11	0,04	0,02	0,01	0,00
Y	ε_π	0,18	0,16	0,11	0,05	0,04	0,02	0,02
	ε_R	0,00	0,00	0,01	0,01	0,01	0,00	0,00
	ε_Y	0,81	0,84	0,88	0,94	0,96	0,97	0,98

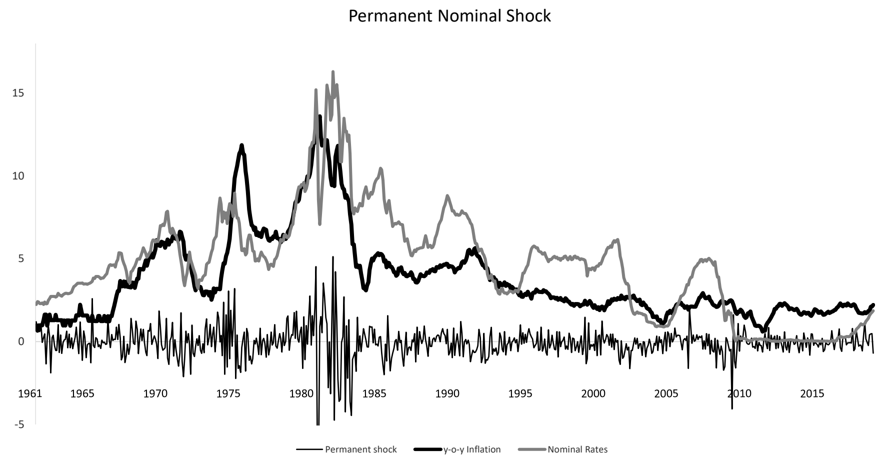
Forecast Error Variance decomposition in VECM model. Estimation samples for France: 1961-2018, or 1961 M8 - 2018 M12 (T=689), ii) 1961-1999, or 1961 M8 - 1999 M12 (T=461) and iii) 1999-2018, or 1999 M1 -2018 M12 (T=240)

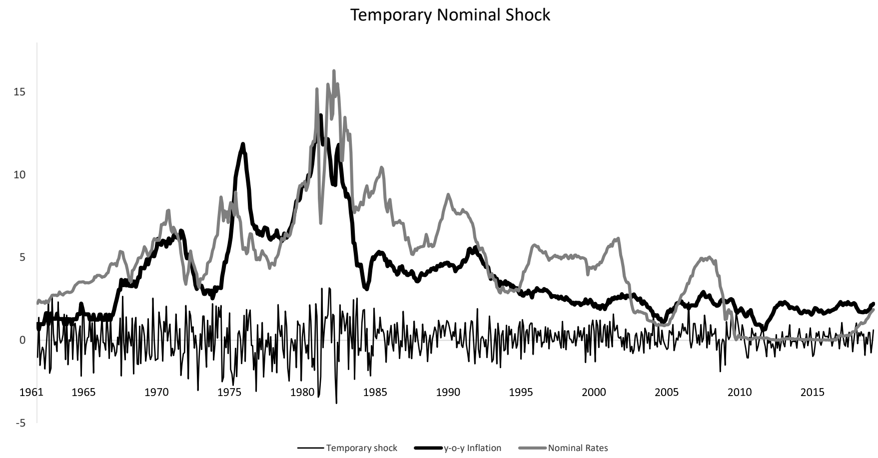
Table 3. Forecast Error Variance Decomposition - France

2.3. Structural Shocks

We present here the identified structural nominal shocks: the permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Shocks are identified as described in the text and are such that $u_t = B\varepsilon_t$, where $\varepsilon_t = (\varepsilon_{\pi,t}, \varepsilon_{R,t}, \varepsilon_{Y,t})'$ is a vector of serially and mutually uncorrelated structural shocks with normalized (unit) variance and $u_t = (u_{\pi,t}, u_{R,t}, u_{Y,t})'$ is the vector of reduced form innovations. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. For each country we report inflation and nominal interest rate data along with the shocks, looking at shocks identified using the full sample or identified using data for each of the subsamples. The IT break date is indicated with a vertical line. At least for the sizeable shocks identified, we find no striking qualitative differences between what obtains in the full sample and in the subsamples, whereas the magnitudes differ often somewhat. We focus our observations on the full sample identification.

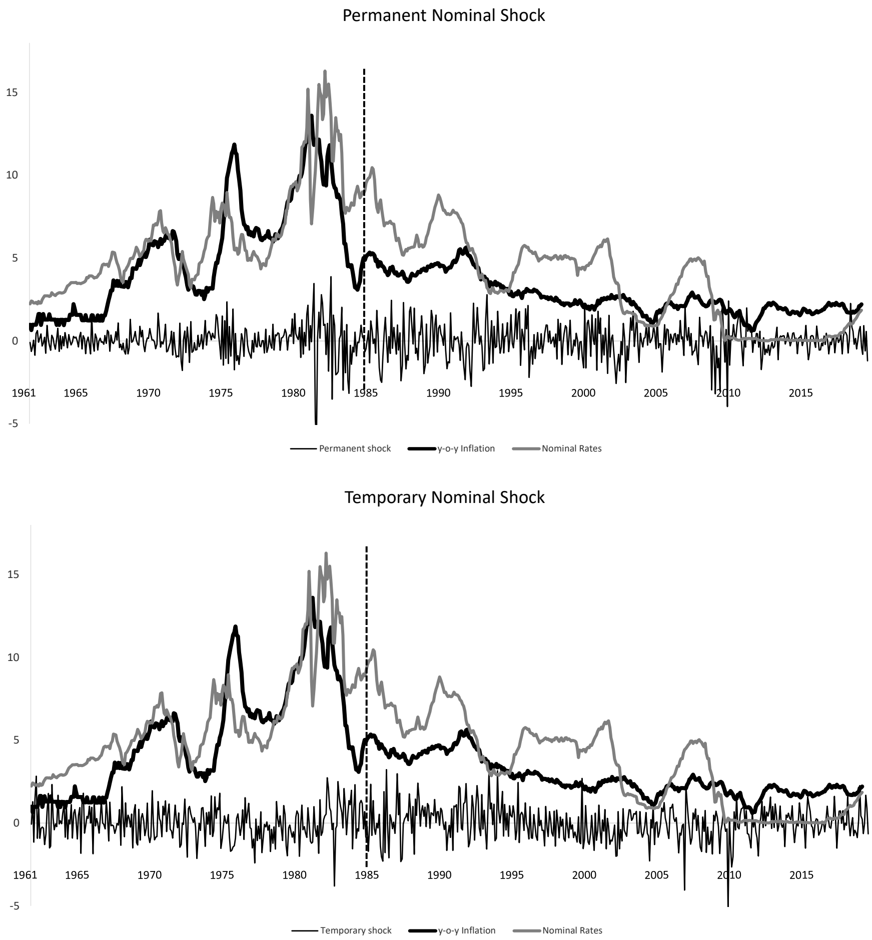
The persistent shifts in inflation and nominal rates in the earlier part of the sample are clearly attributable to the identified permanent shocks (disinflation in the early to mid 1980's in the U.S. and France is attributed to large permanent and negative shocks), although relevant temporary shocks are also present (quite noticeably when spikes in nominal rates are quickly reversed). Permanent shocks also play a relevant role over the more recent period. Interestingly, the gradual shift associated with hitting the effective lower bound in the U.S., Japan and to a lesser extent in France, is attributable to permanent negative shocks (although large temporary shocks are also observed).





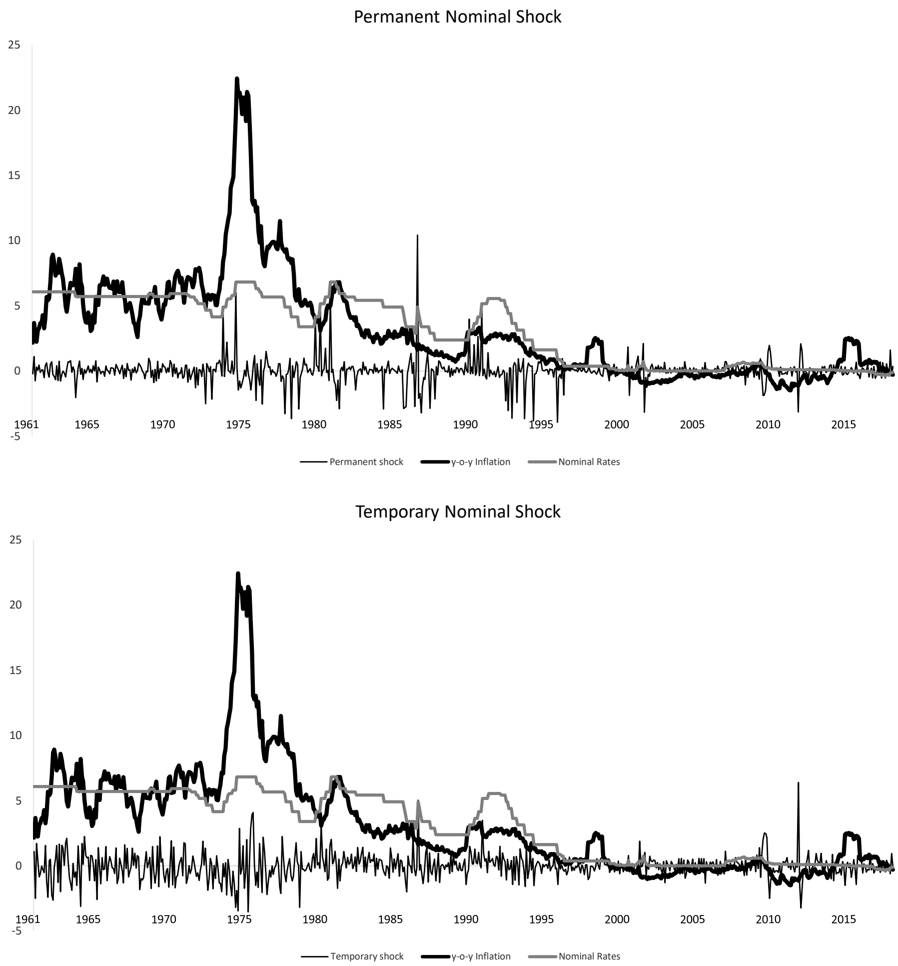
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 4
Identified Nominal Shocks - U.S., Full Sample



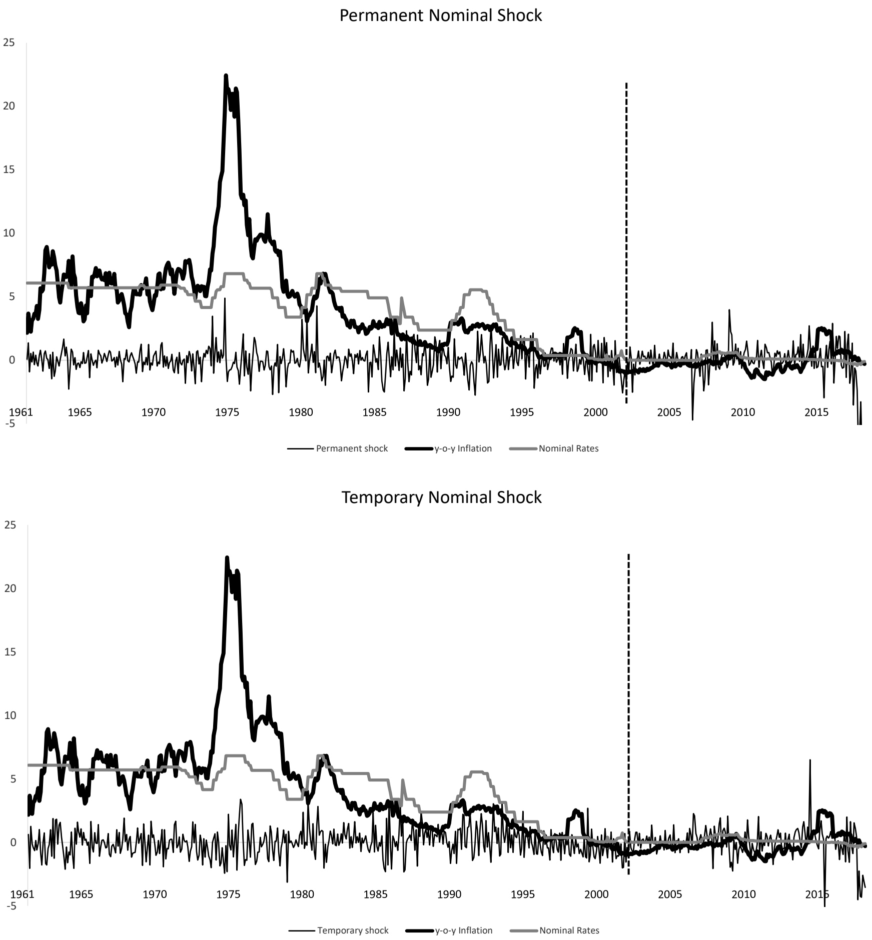
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance). Here the structural shocks are based on the model estimated in each of the sub-samples, Pre-IT and Post-IT.

Figure 5
Identified Nominal Shocks - U.S., using Pre-IT and Post-IT sub-samples



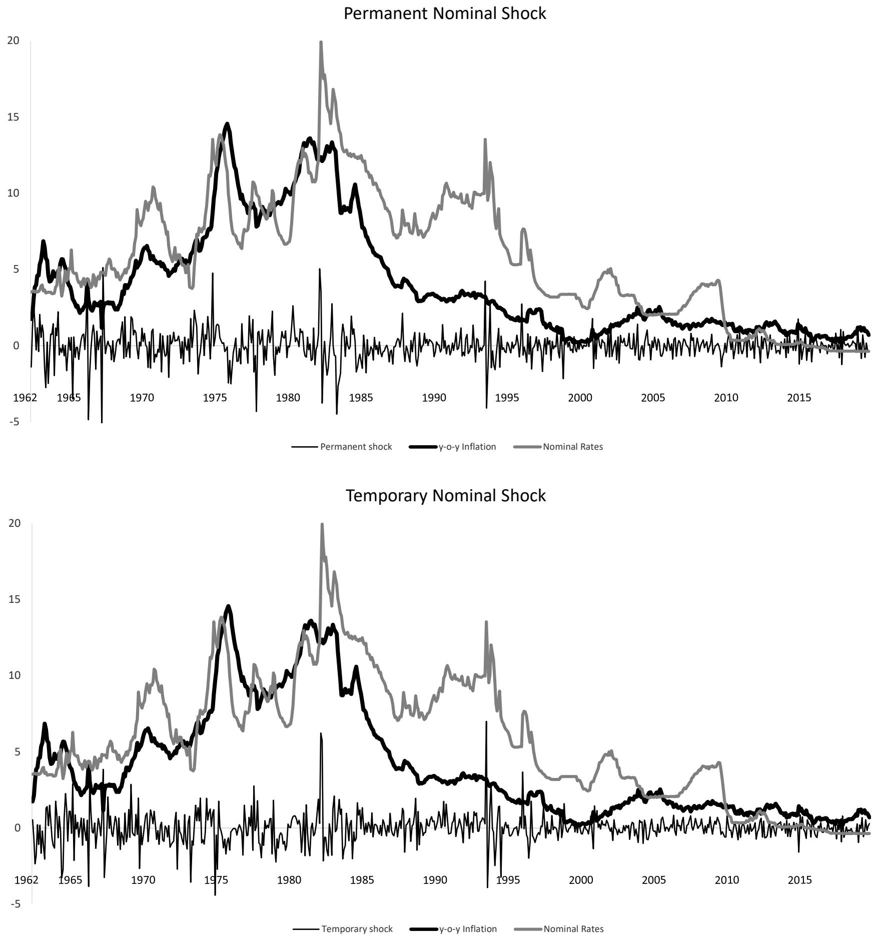
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 6
Identified Nominal Shocks - Japan, Full Sample



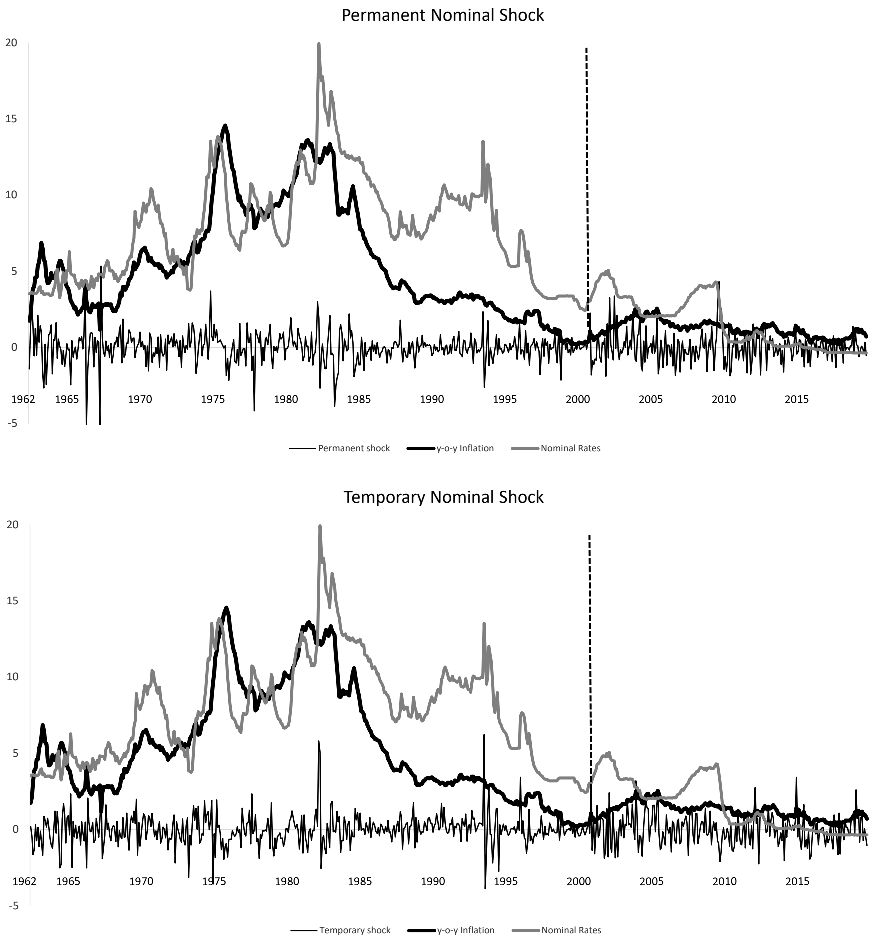
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance). Here the structural shocks are based on the model estimated in each of the sub-samples, Pre-IT and Post-IT.

Figure 7
Identified Nominal Shocks - Japan, using Pre-IT and Post-IT sub-samples



Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 8
Identified Nominal Shocks - France, Full Sample



Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance). Here the structural shocks are based on the model estimated in each of the sub-samples, Pre-IT and Post-IT.

Figure 9
Identified Nominal Shocks - France, using Pre-IT and Post-IT sub-samples

3. Long-run and short-run dynamics in the U.K., Germany and the euro area

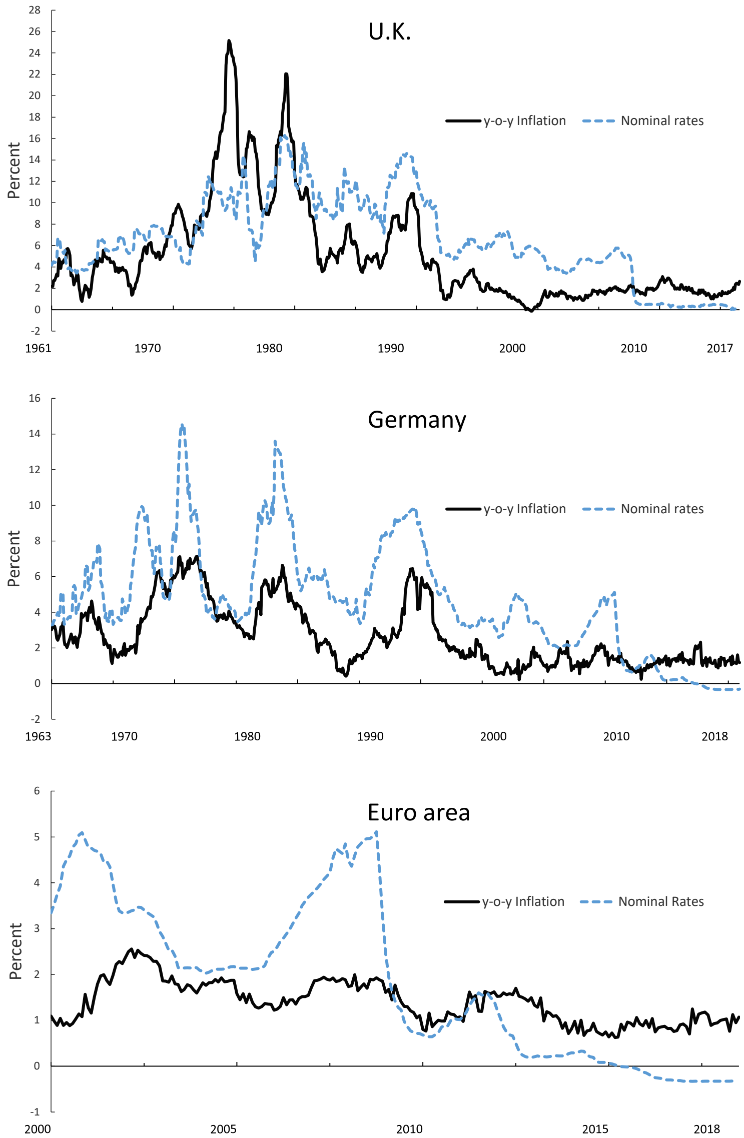


Figure 10
Data - Inflation and Nominal Rates - monthly data for the U.K., Germany and the euro area

We repeat the exercise in the main text and in the first part of this appendix for the U.K., Germany and the euro area. We take again monthly data for the levels of core inflation, nominal interest rates (see Figure 10), and industrial production (our measure of output) for the U.K. (1961-2017), Germany (1963-2018) and the Euro area (2001-2018) and investigate the dynamics of these variables. Again, standard tests reveal that cointegration between inflation and nominal rates is a reasonable hypothesis if one takes the whole sample (and the two variables appear to have a unit-root, even in the case of the euro area). A unit-root also characterizes output.

Table 4 reports the estimation results regarding the cointegration vector and the adjustment parameters, taking the whole sample and also the pre-IT and post-IT subsamples for the U.K. and Germany, where the break date is 1992 for the U.K. and the 1999 (beginning of the euro) for Germany. In the case of the U.K. we note a cointegrating parameter close to 1 for the whole sample (0,84), even closer for the pre-IT period (0,94) and substantially lower value for the post-IT period. For Germany the coefficient is much for the whole sample (0,52). and for the pre-IT sample (0,68) but becomes hardly significant in the post-IT subsample, which questions the reasonableness of the VECM specification for this more stationary period for inflation and nominal rates (just as in the case of Japan). For the euro area, where data is used from 2001 onwards the cointegration coefficient is rather low (0,31) but significant, resembling the behavior for France. Indeed, as we noticed above, a unit-root is a reasonable characterization for inflation and nominal rates in the euro area, even though we observe essentially two not so markedly different regimes: one with relatively high nominal rates and inflation hovering around 2 percent and one with low nominal rates and inflation hovering around 1 percent. That is, we observe shifhts in the two variables but ex-post real interest rates fall relevantly, which justifies the low cointegration coefficient. Again, we notice further that the variable adjusting more significantly towards the cointegration relation is inflation (larger and significant coefficient in most instances).

Turning now to the impulse response functions of the identified structural shocks for the U.K., Germany and the euro area (Figures 11, 12 and 13 for the whole sample and Figures 14, 15 for the pre-IT sample for the U.K. and Germany) we note that the results are strikingly similar to what obtained for the U.S., Japan and France in the case of nominal shocks. Starting with the full sample and with the permanent output shock, we notice that it is associated with temporarily lower inflation in all cases but temporarily lower nominal rates in the case of Germany and the euro area (in the U.K nominal rates rise). We notice that in the euro area the effects on inflation are small. The effects of the temporary nominal shock are again the conventional ones: a temporary shock increases nominal rates and drives inflation downwards. Output falls on impact in the U.K. and Germany, although it recovers quickly. For the euro area no contractionary effects are found, although they are small. All these impacts of the temporary nominal shock essentially fade away after three years, closer to four years in the case of inflation in Germany. We also notice that in the three cases the moderate increase in nominal rates is associated with a somewhat stronger fall in inflation. Turning now to the permanent

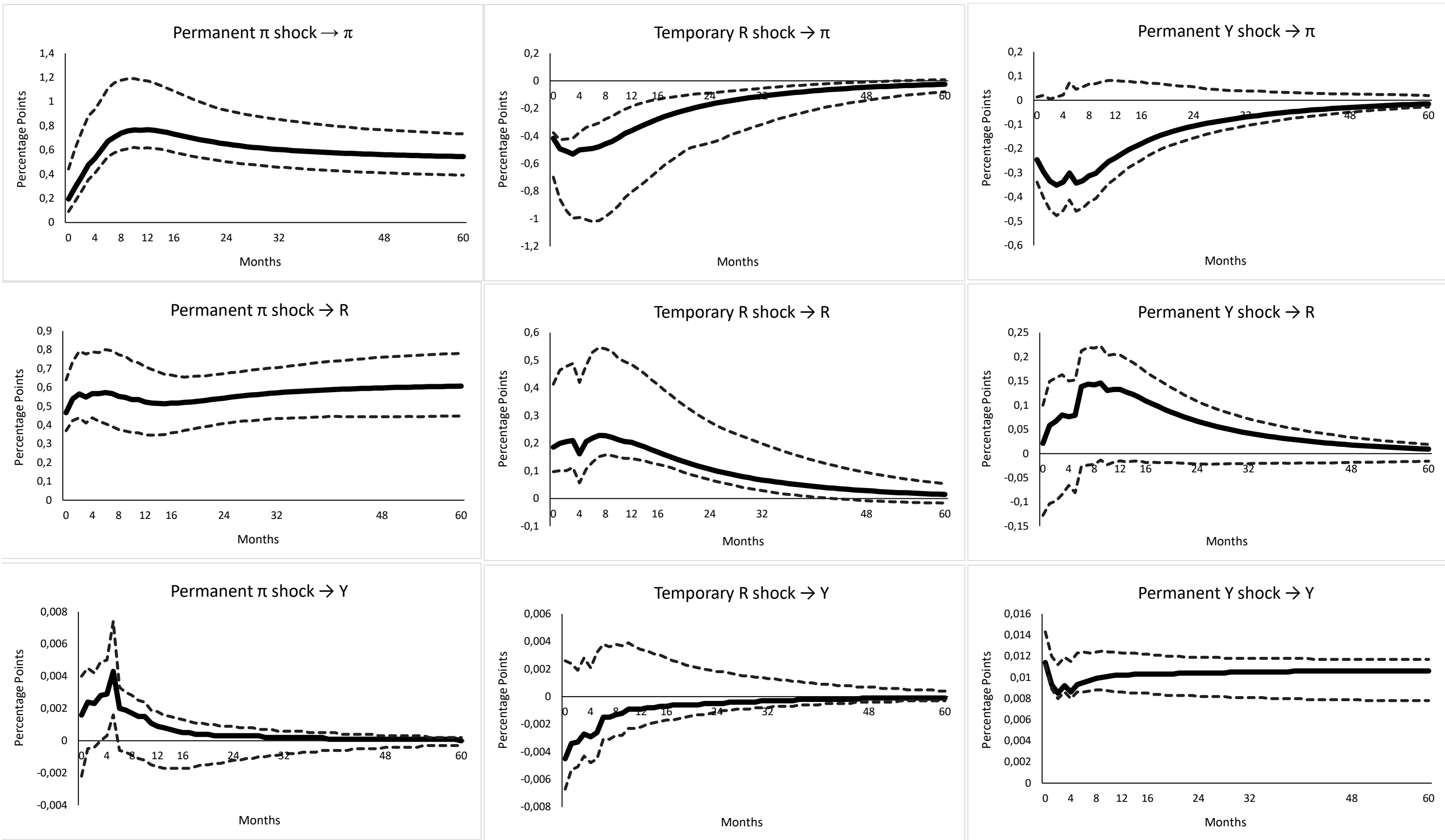
Estimation of Cointegration parameters δ, $\gamma := (\gamma_\pi, \gamma_R, \gamma_Y)$												
Sample	U.K.				Germany				Euro area (2001-)			
	δ	γ_π	γ_R	γ_Y	δ	γ_π	γ_R	γ_Y	δ	γ_π	γ_R	γ_Y
1961*-2018	0,86	-0,02	0,01	0,000	0,52	-0,04	0,016	0,000	0,31	-0,12	0,003	0,001
	0,18	0,01	0,001	0,000	0,08	0,01	0,013	0,001	0,05	0,03	0,024	0,002
1961*-IT	0,94	-0,03	0,02	0,000	0,68	-0,05	0,037	-0.001	-	-	-	-
	0,33	0,001	0,001	0,000	0,11	0,01	0,02	0.001				
IT-2018	0,08	-0,1	0.015	-0,002	-0,01	-0,22	0.013	0,001	-	-	-	-
	0,04	0,02	0,02	0,001	0,04	0,05	0,02	0,003				

Estimates of cointegration parameters in the VECM model. Standard errors are below the estimates. Cointegration rank is restricted to 1 and coefficient of output in the cointegration relation is restricted to zero. 6 lags of differenced endogenous variables are considered in the VECM. Simple two step estimator (S2S), as implemented in JMulTi is employed. Estimation samples for the U.K: i) 1961*-2018, or 1961 M8 - 2017 M6 (T=671), ii) 1961*-IT, or 1961 M8 -1992M12 (T=377) and iii) IT-2018, or 1992 M1 -2017 M6 (T=306); Estimation samples for Germany: i) 1961*-2018, or 1963 M8 - 2018 M12 (T=665), ii) 1961*-IT, or 1963 M8 - 1999 M12 (T=437) and iii) IT-2018, or 1999 M1 -2018 M12 (T=240); Estimation samples for the Euro area 2001 M1 -2018 M5 (T=209).

Table 4. Vector Error Correction model Estimation. Cointegration parameters

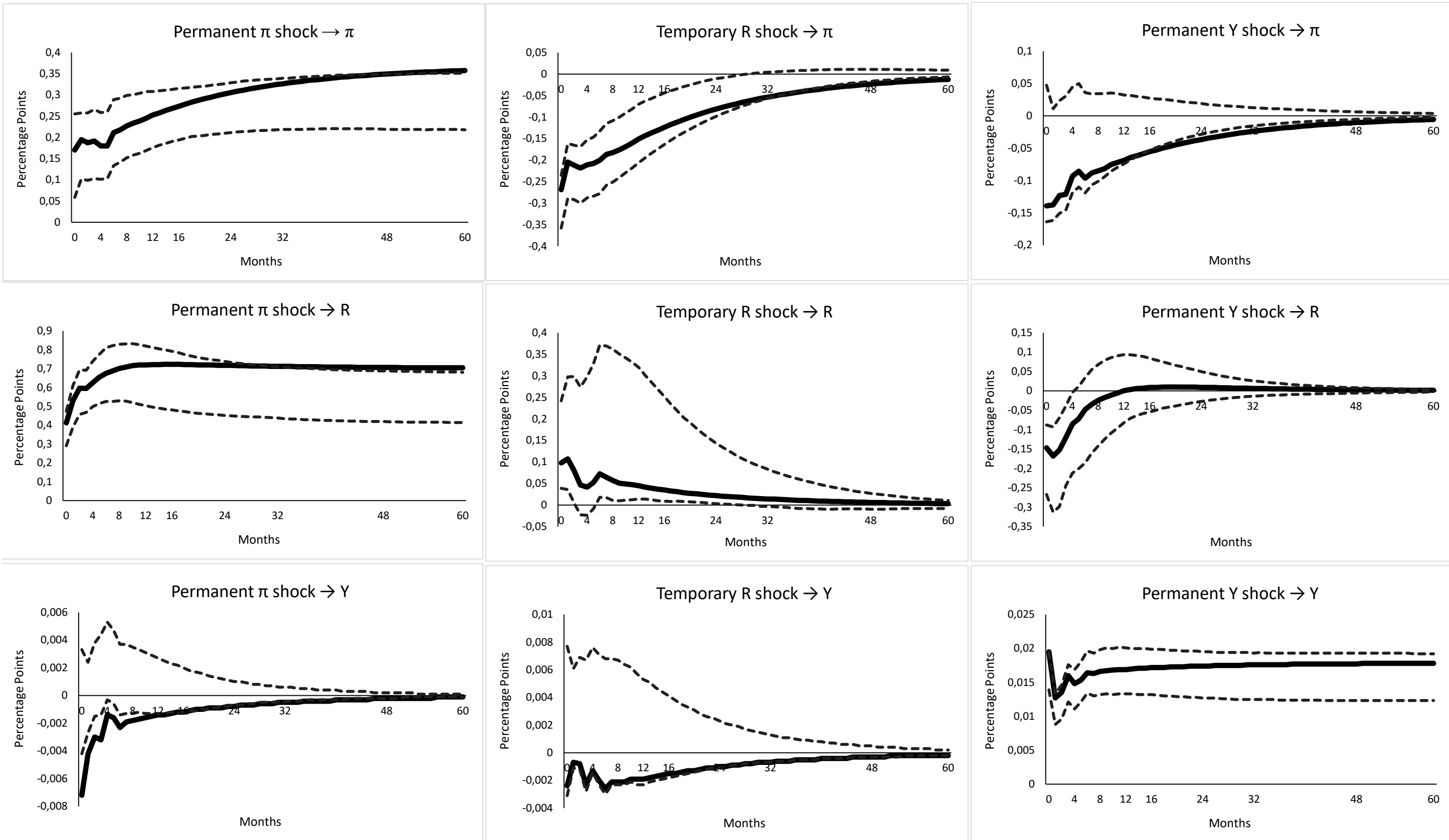
nominal shock (leading to a permanent increase in nominal rates) we notice that, as expected given the cointegration parameters, the permanent increase in nominal rates is accompanied by an increase in inflation towards the new long-run level. What we highlight again is that this adjustment takes place in roughly less than two years in all cases while qualitatively the short-run effect of this permanent shock is very similar to the long-run one. No fall of inflation in the short run is observed, i.e., we find again the Neo-Fisher effect first documented by Uribe (2017, 2018). The effects of this permanent nominal shock are contractionary for Germany and the euro area but not for the U.K.

Finally, the characterization above fits almost always very well to what obtains in the pre-IT sample for the U.K. and Germany.



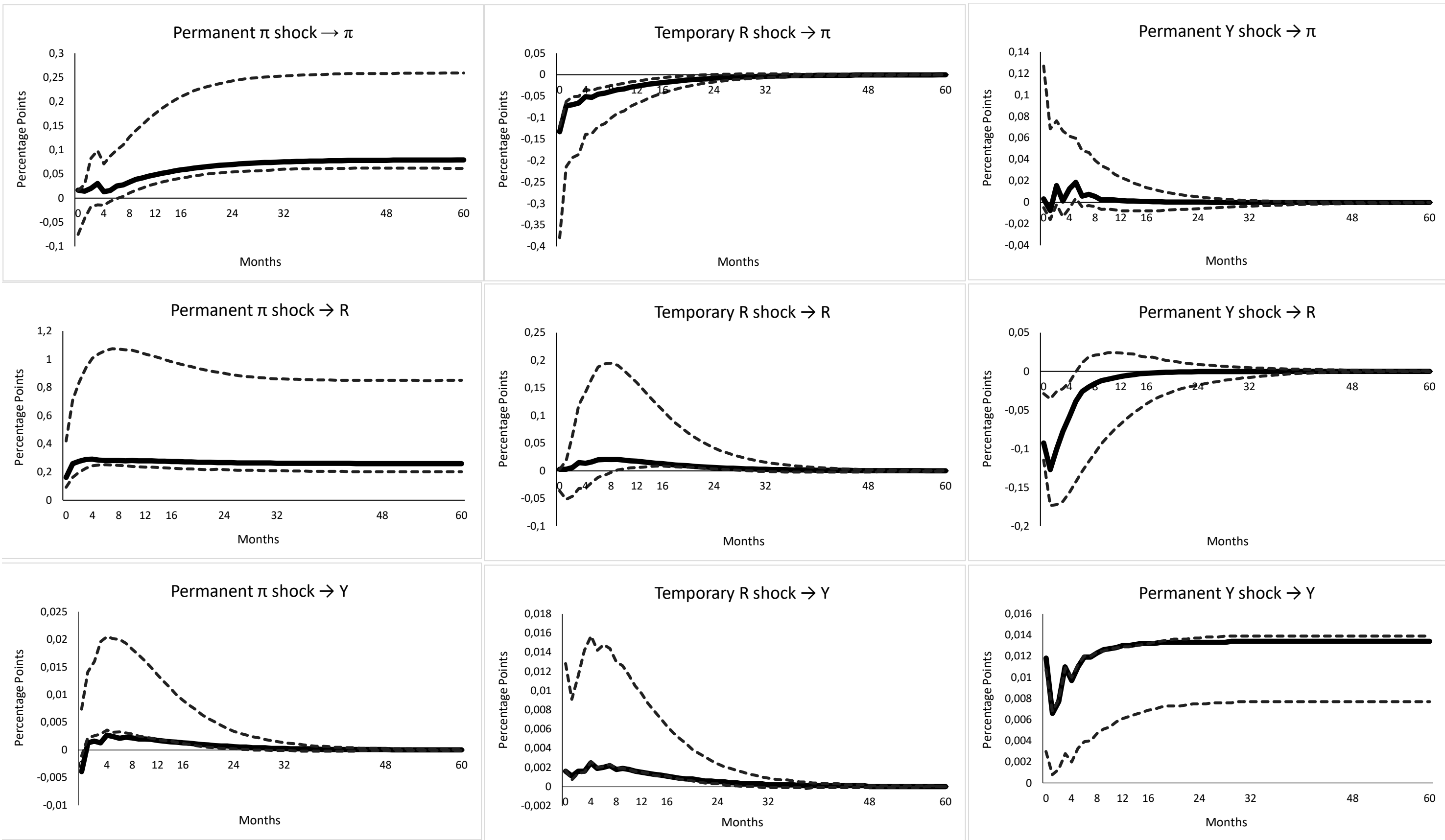
Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 11
Impulse Response Functions - U.K., Full Sample



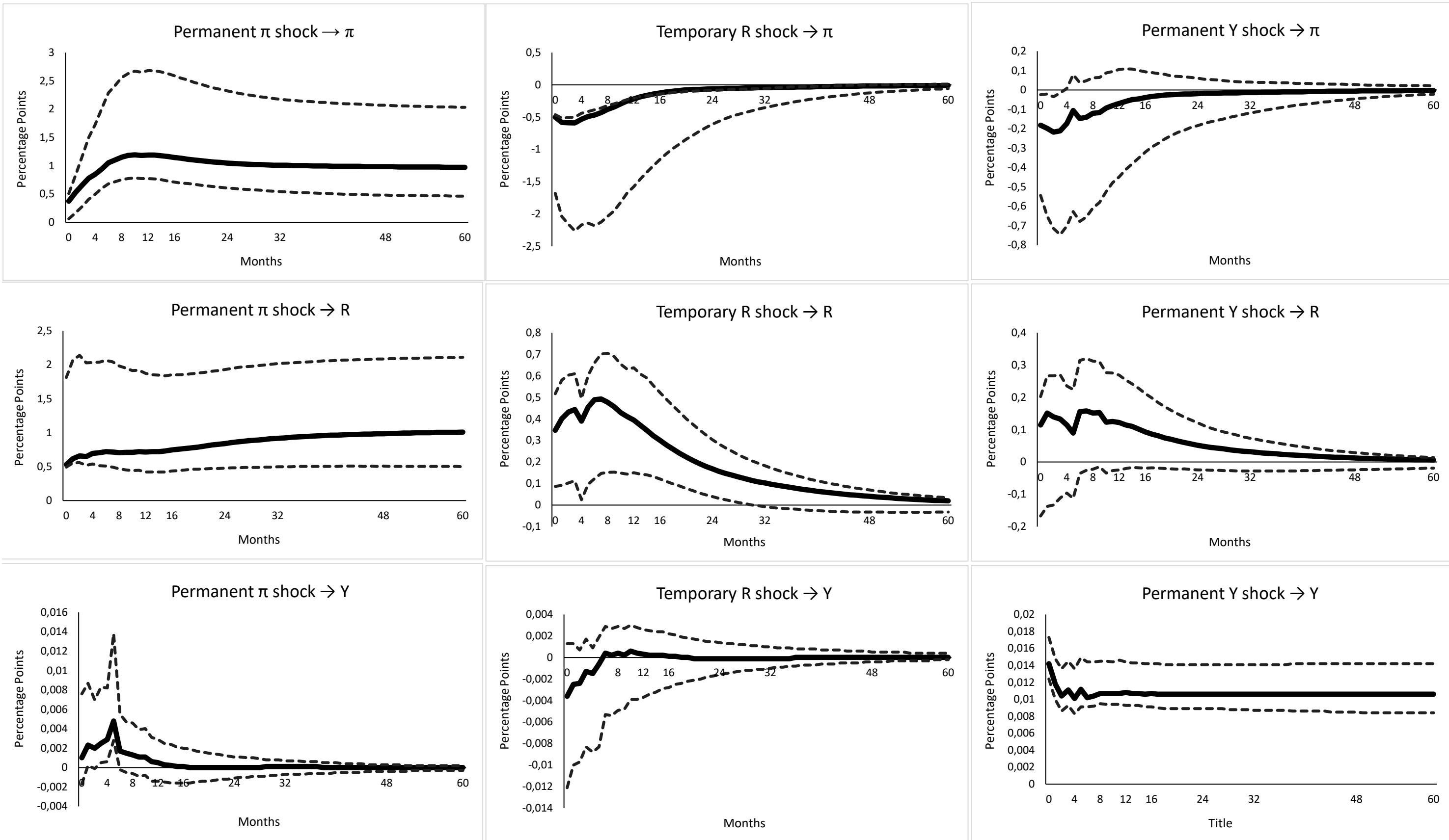
Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 12
Impulse Response Functions - Germany, Full Sample



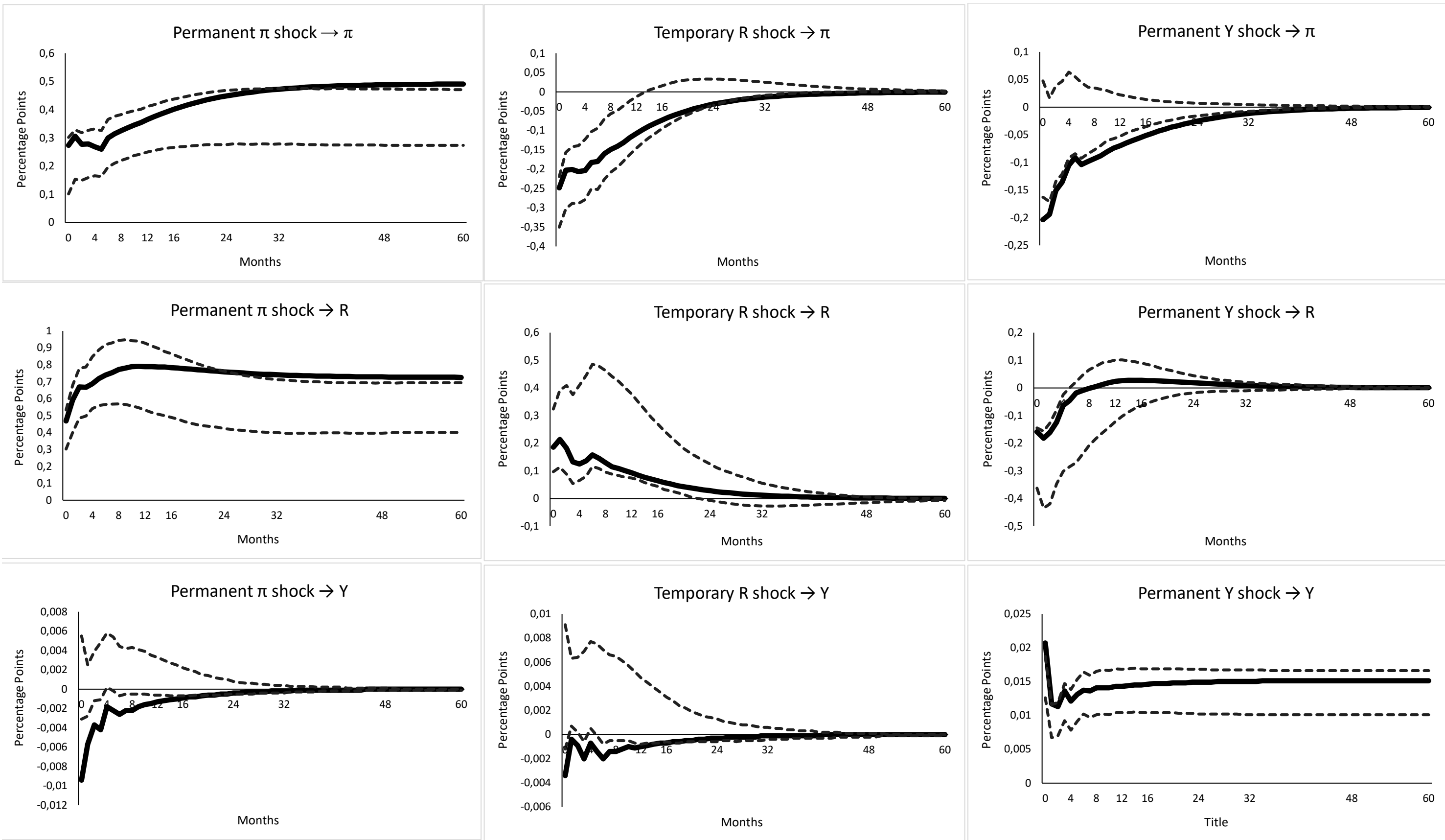
Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 13
Impulse Response Functions - Euro area, Full Sample



Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 14
Impulse Response Functions - U.K., Pre-IT Sample



Response of inflation, nominal rates and output to identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$), temporary nominal shock ($\varepsilon_{R,t}$)and permanent output shock ($\varepsilon_{Y,t}$). Shocks are identified as described in the text. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. Response to a one standard deviation shock, in percentage points. 90 % Hall Bootstrap confidence intervals.

Figure 15
Impulse Response Functions - Germany, Pre-IT Sample

3.1. Forecast Error Variance Decomposition

Next we analyze the decomposition of the variance of the forecast error for π_t, R_t and Y_t. Tables 5, 6 and 7 display the results for the three economies. We find several similarities with the analysis made above for the U.S. and France, with the U.K. and Germany displaying similar patterns and resembling the U.S. characterization when one takes the full sample or the pre-IT sample. First, in all cases we find again that the bulk of the variance of the forecast error in the case of output is attributed to the output shock, even at short horizons. Also, this shock also accounts for some fraction of the variance of the forecast error of inflation in several instances, but much less so for the U.K. if one takes the post-IT sample and for the euro area (at least for the U.K. in the post-IT sample, results ought to be read with caution due to potential stationarity of nominal rates and inflation and imprecision in the estimation). Further, the contribution of the permanent nominal shock for the variance of the forecast error of nominal rates is again always very high, even at short horizons. At short horizons, this shock also accounts for a relevant fraction of the variance of the forecast error of inflation (much less so in the case of Germany for the post-IT sample). As for the temporary nominal shock, it often contributes relevantly to the variance of the forecast error of inflation at short horizons, and more so in the post-IT sample for Germany and for the euro area. It also contributes somewhat to the variance of the forecast error of nominal rates (but less so in the euro area and in the post-IT sample for Germany). Again, the contribution of this temporary shock to the variance of the forecast error of output is generally rather low, except in a few instances at short horizons.

UK 1961-2017

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,14	0,18	0,27	0,54	0,68	0,78	0,81
	ε_R	0,64	0,61	0,52	0,32	0,23	0,15	0,13
	ε_Y	0,22	0,22	0,21	0,14	0,10	0,07	0,06
R	ε_π	0,86	0,87	0,86	0,84	0,86	0,92	0,94
	ε_R	0,14	0,13	0,12	0,12	0,10	0,06	0,05
	ε_Y	0,00	0,01	0,01	0,03	0,03	0,02	0,00
Y	ε_π	0,02	0,03	0,05	0,05	0,03	0,01	0,01
	ε_R	0,13	0,12	0,11	0,06	0,03	0,02	0,01
	ε_Y	0,85	0,84	0,84	0,89	0,94	0,97	0,98

UK 1961-1992

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,33	0,38	0,50	0,79	0,89	0,94	0,95
	ε_R	0,60	0,55	0,44	0,19	0,10	0,05	0,04
	ε_Y	0,08	0,07	0,06	0,02	0,01	0,01	0,00
R	ε_π	0,68	0,68	0,67	0,69	0,78	0,90	0,92
	ε_R	0,29	0,29	0,29	0,28	0,20	0,09	0,07
	ε_Y	0,03	0,04	0,03	0,03	0,02	0,01	0,00
Y	ε_π	0,00	0,02	0,03	0,04	0,02	0,01	0,01
	ε_R	0,06	0,05	0,04	0,02	0,01	0,01	0,00
	ε_Y	0,94	0,93	0,93	0,94	0,97	0,98	0,99

UK 1992-2017

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,10	0,15	0,23	0,37	0,42	0,44	0,45
	ε_R	0,57	0,52	0,45	0,28	0,25	0,24	0,24
	ε_Y	0,33	0,33	0,32	0,35	0,33	0,32	0,31
R	ε_π	0,94	0,95	0,95	0,97	0,98	0,99	0,99
	ε_R	0,01	0,01	0,00	0,01	0,01	0,00	0,00
	ε_Y	0,04	0,05	0,05	0,02	0,01	0,01	0,00
Y	ε_π	0,00	0,00	0,08	0,10	0,04	0,02	0,01
	ε_R	0,19	0,16	0,14	0,07	0,03	0,01	0,01
	ε_Y	0,81	0,83	0,77	0,84	0,93	0,97	0,98

Forecast Error Variance decomposition in VECM model. Estimation samples for the U.K: i) 1961-2017, or 1961 M8 - 2017 M6 (T=671), ii) 1961*-1992, or 1961 M8 -1992 M12 (T=377) and iii) 1992-2017, or 1992 M1 -2017 M6 (T=306)

Table 5. Forecast Error Variance Decomposition - U.K.

Germany 1963-2018

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,24	0,31	0,34	0,46	0,64	0,82	0,86
	ε_R	0,60	0,52	0,50	0,43	0,29	0,14	0,11
	ε_Y	0,16	0,17	0,17	0,11	0,07	0,03	0,03
R	ε_π	0,84	0,86	0,91	0,97	0,98	0,99	0,99
	ε_R	0,05	0,04	0,02	0,01	0,01	0,00	0,00
	ε_Y	0,11	0,10	0,07	0,02	0,01	0,00	0,00
Y	ε_π	0,12	0,11	0,08	0,03	0,02	0,01	0,01
	ε_R	0,01	0,01	0,01	0,01	0,01	0,01	0,00
	ε_Y	0,87	0,88	0,91	0,95	0,97	0,99	0,99

Germany 1963-1999

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,42	0,48	0,51	0,66	0,82	0,93	0,94
	ε_R	0,35	0,29	0,30	0,23	0,12	0,05	0,04
	ε_Y	0,23	0,23	0,19	0,11	0,06	0,02	0,02
R	ε_π	0,79	0,80	0,86	0,94	0,97	0,98	0,99
	ε_R	0,12	0,11	0,08	0,04	0,02	0,01	0,01
	ε_Y	0,09	0,08	0,06	0,02	0,01	0,00	0,00
Y	ε_π	0,17	0,17	0,15	0,07	0,04	0,02	0,01
	ε_R	0,02	0,02	0,02	0,01	0,01	0,00	0,00
	ε_Y	0,81	0,81	0,84	0,92	0,96	0,98	0,98

Germany 1999-2018

Var.	Shock	Horizon 1	2	4	12	24	48	60
π	ε_π	0,04	0,10	0,16	0,18	0,20	0,20	0,21
	ε_R	0,96	0,85	0,73	0,72	0,70	0,70	0,69
	ε_Y	0,00	0,05	0,11	0,10	0,10	0,10	0,10
R	ε_π	0,75	0,79	0,85	0,94	0,98	0,99	0,99
	ε_R	0,00	0,00	0,00	0,00	0,00	0,00	0,00
	ε_Y	0,25	0,21	0,14	0,06	0,02	0,01	0,00
Y	ε_π	0,32	0,29	0,22	0,09	0,04	0,02	0,02
	ε_R	0,00	0,00	0,00	0,00	0,00	0,00	0,00
	ε_Y	0,68	0,71	0,78	0,91	0,95	0,98	0,98

Forecast Error Variance decomposition in VECM model. Estimation samples for Germany: i) 1963*-2018, or 1963 M8 - 2018 M12 (T=665), ii) 1961*-IT, or 1963 M8 - 1991 M12 (T=437) and iii) IT-2018, or 1999 M1 -2018 M12 (T=240)

Table 6. Forecast Error Variance Decomposition - Germany

Euro area 2001-2018

Var.	Shock	Horizon						
		1	2	4	12	24	48	60
π	ε_π	0.02	0.02	0.05	0.18	0.52	0.79	0.84
	ε_R	0.98	0.98	0.94	0.81	0.48	0.21	0.16
	ε_Y	0.00	0.01	0.01	0.02	0.01	0.00	0.00
R	ε_π	0.75	0.79	0.86	0.95	0.97	0.98	0.99
	ε_R	0.00	0.00	0.00	0.00	0.00	0.00	0.00
	ε_Y	0.25	0.21	0.14	0.05	0.03	0.01	0.01
Y	ε_π	0.10	0.08	0.05	0.04	0.02	0.01	0.01
	ε_R	0.02	0.02	0.02	0.03	0.01	0.01	0.01
	ε_Y	0.89	0.90	0.92	0.94	0.97	0.98	0.99

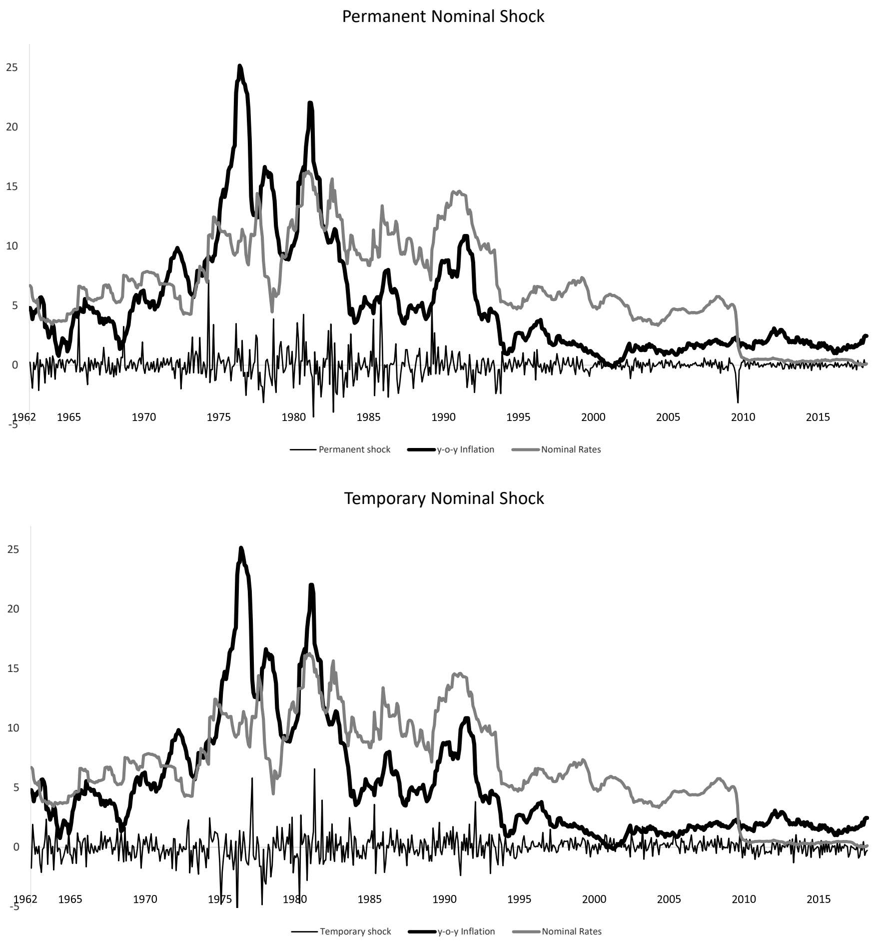
Forecast Error Variance decomposition from VECM model. Estimation sample for the Euro area 2001 M1 - 2018 M5 (T=209).

Table 7. Forecast Error Variance Decomposition - euro area

3.2. Structural Shocks

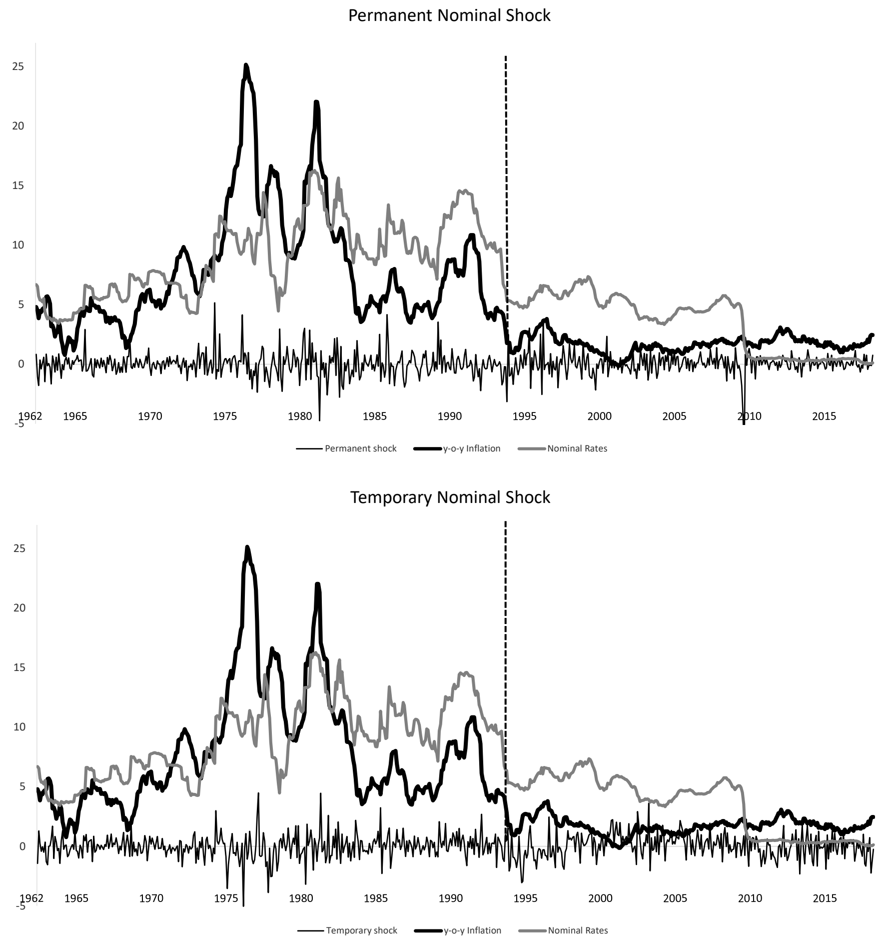
We present here the identified structural nominal shocks for the U.K., Germany and the euro area: the permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Shocks are identified as described in the text and are such that $u_t = B\varepsilon_t$, where $\varepsilon_t = (\varepsilon_{\pi,t}, \varepsilon_{R,t}, \varepsilon_{Y,t})'$ is a vector of serially and mutually uncorrelated structural shocks with normalized (unit) variance and $u_t = (u_{\pi,t}, u_{R,t}, u_{Y,t})'$ is the vector of reduced form innovations. The structural model is estimated by Maximum Likelihood, using the Scoring Algorithm of Amisano and Giannini (1997) as implemented in JMulTi. For each country we report inflation and nominal interest rate data along with the shocks, looking at shocks identified using the full sample or identified using data for each of the subsamples (no subsamples for the euro area). The IT break date is indicated with a vertical line. We stress again that given the normalization of the variance, the analysis of these plots ought to be complemented with the results for the forecast error variance decomposition. E.g., large identified temporary shocks may contribute little to the dynamics of nominal rates or inflation if their contribution to the variance of the forecast errors is small.

We focus our observations on the full sample identification. We notice again that at least for the sizeable shocks identified, we find no striking qualitative differences between what obtains in the full sample and in the subsamples. The magnitudes differ often somewhat. We find again clear similarities to what obtained for the U.S., France and Japan. For the U.K. and Germany, the persistent shifts in inflation and nominal rates in the earlier part of the sample are again clearly attributable to the identified permanent shocks, although the size of the two shocks is more similar in the case of the U.K.. Also, the observed disinflation in the early to mid 1980's in the U.K. and Germany can be attributed to large permanent and negative shocks. Relevant temporary shocks are also present in the earlier part of the sample. We also notice that the volatility of the permanent shocks tends to be lower in the later part of the sample. The same is true for the temporary shocks but only the case of the U.K. Permanent shocks play again a relevant role over the more recent period. The recent effective lower bound episode in the U.K., Germany and also the euro area, is attributable to a large permanent negative shock (although somewhat large temporary shocks are also observed during this period).



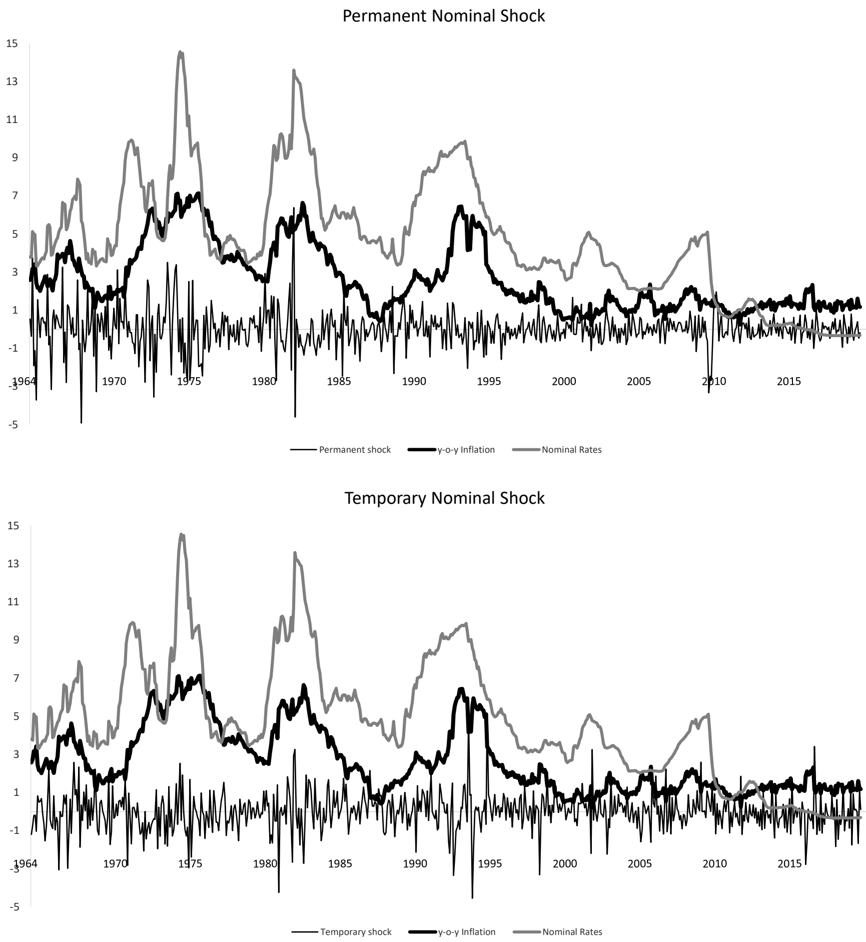
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 16
Identified Nominal Shocks - U.K. Full, Sample



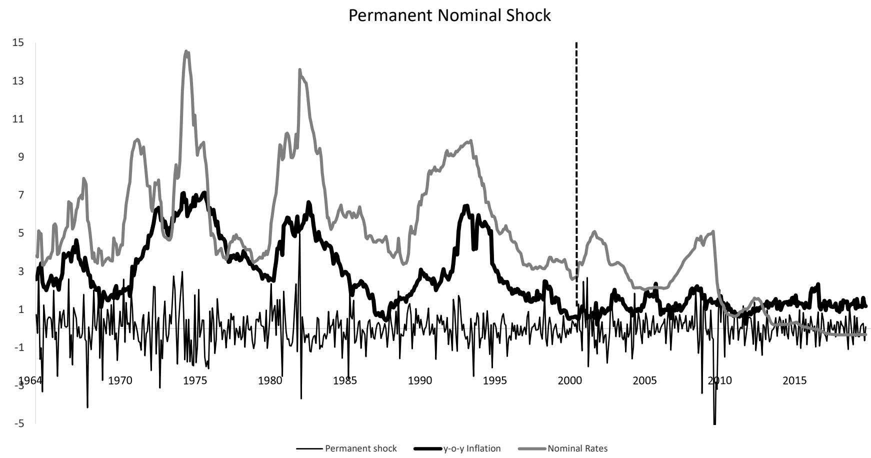
Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance). Here the structural shocks are based on the model estimated in each of the sub-samples, Pre-IT and Post-IT.

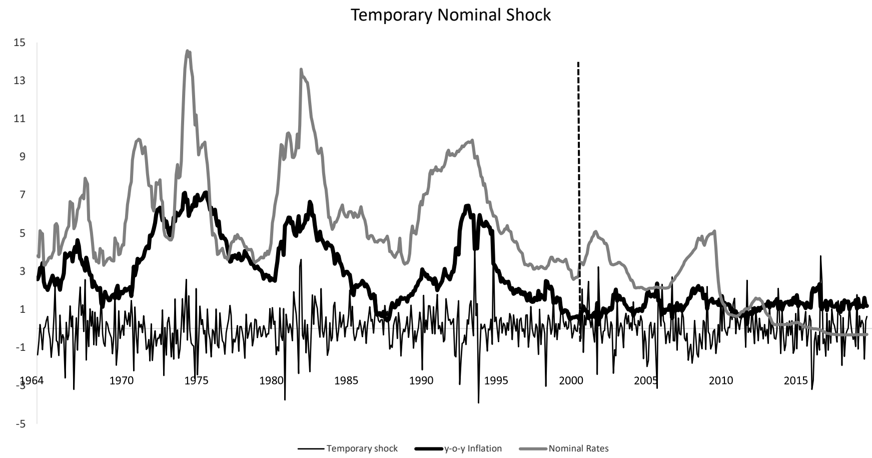
Figure 17
Identified Nominal Shocks - U.K., using Pre-IT and Post-IT sub-samples



Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

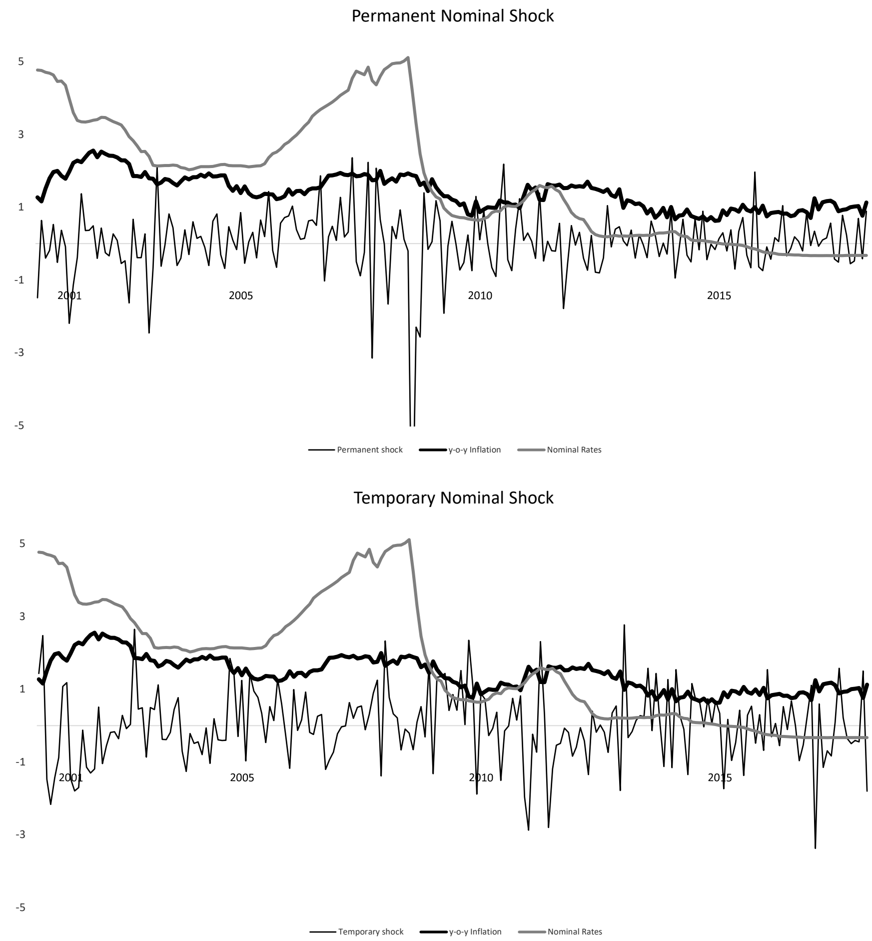
Figure 18
Identified Nominal Shocks - Germany, Full Sample





Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance). Here the structural shocks are based on the model estimated in each of the sub-samples, Pre-IT and Post-IT.

Figure 19
Identified Nominal Shocks - Germany, using Pre-IT and Post-IT sub-samples



Identified structural shocks: permanent nominal shock ($\varepsilon_{\pi,t}$, Top Panel) and temporary nominal shock ($\varepsilon_{R,t}$, Bottom panel). Structural shocks are normalized (unit variance).

Figure 20
Identified Nominal Shocks - Euro area, Full Sample

Description	FRED code
Consumer Price Index for All Urban Consumers: All Items Less Food and Energy	CPILFESL
Consumer Price Index: OECD Groups: All Items Non-Food and Non-Energy (Japan)	CPGRLE01JPM659N
Harmonized Index of Consumer Prices: Overall Index Excluding Energy, Food, Alcohol, and Tobacco for Euro area (EA11-2000, EA12-2006, EA13-2007, EA15-2008, EA16-2010, EA17-2013, EA18-2014, EA19)	00XEFDEZCCM086NEST
Consumer Price Index of All Items in France (only up to 1970 M12)	FRACPIALLMINMEI
Consumer Price Index: All Items Excluding Food and Energy for France (from 1971 M1 onwards)	FRACPICORMINMEI
Consumer Price Index of All Items in the United Kingdom (only up to 1970 M12)	GBRCPIALLMINMEI
Consumer Price Index: All Items Excluding Food and Energy for U.K. (from 1971 M1 onwards)	GBRCPICORMINMEI
Consumer Price Index: All Items Excluding Food and Energy for Germany	DEUCPICORMINMEI
Industrial Production Index (U.S.)	INDPRO
Production of Total Industry in Japan (Industrial Production Index)	JPNPROINDMISMEI
Euro area 19 (fixed composition) - Industrial Production Index, Total Industry - NACE Rev2	[Statistical Data Warehouse -ECB]
Production of Total Industry in France	FRAPROINDMISMEI
Production of Total Industry in the United Kingdom	GBRPROINDMISMEI
Production of Total Industry in Germany	DEUPROINDMISMEI
3-Month Treasury Bill: Secondary Market Rate (U.S.)	TB3MS
Interest Rates, Government Securities, Treasury Bills for Japan	INTGSTJPM193N
3-Month or 90-day Rates and Yields: Interbank Rates for the Euro Area	IR3TIB01EZM156N
Immediate Rates: Less than 24 Hours: Call Money/Interbank Rate for France	IRSTCI01FRM156N
3-Month or 90-day Rates and Yields: Treasury Securities for the U.K.	IR3TTS01GBM156N
3-Month or 90-day Rates and Yields: Interbank Rates for Germany	IR3TIB01DEM156N

Note: sources are Federal Reserve Bank of St. Louis's FRED and ECB's Statistical Data Warehouse

Table 8. Monthly data for core inflation, short rates and industrial production

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