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**Markups' Cyclical Behavior: The role of
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Markups' Cyclical Behavior: The role of Demand and Supply Shocks^{*}

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Abstract

We assess how demand and supply shocks (identified via the Blanchard and Quah (1989) SVAR approach) in 14 OECD countries affect mark-ups. We find that individual responses of markups to demand shocks push down the markup for most countries (confirmed in the panel analysis). On the other hand, a supply shock has a more mixed effect.

JEL: C23, E32, E62

Keywords: Blanchard-Quah, mark-up, VAR, impulse response function, local projection

1. INTRODUCTION

The interaction between fiscal policy effectiveness and imperfect competition has received some attention in economic theory (Hall 2009; Christiano et al. 2011; and Woodford 2011; and the survey by Costa and Dixon, 2011). In particular, the cyclical behavior of markups following government spending shocks has been closely analyzed. The New Keynesian Synthesis have developed models that produce undesired endogenous markups due to nominal rigidity, enhancing the effectiveness of demand-side policy, including fiscal policy. Moreover, macroeconomic models with time-varying desired markups are even more attractive as they work similarly to productivity shocks in the presence of active fiscal policy (Ravn et al., 2006).

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The theoretical literature on endogenous markups is dominated by the view that markups behave counter-cyclically following a demand shock. Indeed, when a positive shock originates in the demand side, the marginal cost function is only indirectly affected and the main effect depends on how the individual demand function responds (see e.g. Gali, 1994; Goodfriend and King, 1997; Clarida et al. 1999; Ravn et al., 2006). However, with a positive supply shock, we expect marginal costs to decrease for a given output. Therefore, assuming that the indirect effect on prices via demand is small, markups tend to increase implying a pro-cyclical average markup.

Several papers that try to measure markups for different industries over a period find evidence of a mildly counter-cyclical behavior (Rotemberg and Woodford, 1999; Martins and Scarpetta, 2002), with Afonso and Costa (2013) reporting that markups are pro-cyclical with productivity and counter-cyclical government spending. The combination of demand and supply shocks is a possible explanation for the existing evidence on counter-cyclicality.

We contribute to this literature by looking at how demand and supply shocks in 14 OECD countries affect the evolution of markups in the short to medium run. The two types of shocks are decomposed via the Blanchard and Quah (1989) (BQ) methodology based on a SVAR identification. We conduct both individual country analyses and a panel assessment.

Our results show that individual responses of the markups to demand shocks show that a demand shock pushes down the markup for most countries (confirmed in panel analysis). On the other hand, a supply shock has a more mixed effect.

2. METHODOLOGY

The market is the price wedge

$$\mu_{it} = \frac{P_{it}}{MC_{it}} \quad (1)$$

where P_{it} represents the price of the good produced by firm i and MC_{it} stands for its marginal cost, in t .

Since marginal costs are not observable, one can estimate them using the relationship $MC_{it} = W_t/MPL_{it}$ where W_t is the nominal wage rate (assuming homogeneous labor input) and MPL_{it} is the marginal product of labor. We draw on the markup related data set computed by Afonso and Costa (2013).

In order to separate demand from supply shocks we follow Blanchard and Quah (1989) approach applied to real GDP and unemployment rate data in quarterly frequency between 1970 and 2007. The resulting quarterly series for the shocks are then converted into annual frequency in order to match the markup series.

Under the Blanchard-Quah decomposition, consider the following bivariate SVAR model:

$$A_0 X_t = A_1(L) X_t + B \varepsilon_t \quad (2)$$

where $X_t = (\Delta y_t, \Delta u_t)$ and y and u are measures of real output and unemployment, respectively.¹ Here $\varepsilon_t = (\varepsilon_t^s, \varepsilon_t^d)$ with ε_t^s and ε_t^d being one standard deviation supply and demand shocks, respectively. A_0 is a 2x2 matrix, $A_1(L) = \sum_{i=1}^q A_{1i} L^i$ shoes matrices of lag coefficients of the SVAR system. The BQ approach assumes that two shocks are not correlated, and hence, B is a diagonal matrix. Denote the diagonal elements (essentially are the standard deviations of the two shocks) in B by b_{11} and b_{12} .

The structural shocks in Equation (1) are not directly observable. It is the usual practice to estimate the reduced form VAR and use the estimated parameters and residuals in the reduced form VAR to retrieve the structural shocks. The reduced form VAR has the form:

$$X_t = C(L) X_t + e_t \quad (3)$$

with $C(L) = \sum_{i=1}^q C_i L^i$ being the matrices of estimated lag coefficients and e_t being the vector of two residual series. Equivalently, this reduced form VAR can be expressed in more simple way as:

$$\begin{aligned} \Delta y_t &= \sum_{i=1}^q c_{11}^i \Delta y_{t-1} + \sum_{i=1}^q c_{12}^i \Delta u_{t-1} + e_t^y \\ \Delta u_t &= \sum_{i=1}^q c_{21}^i \Delta u_{t-1} + \sum_{i=1}^q c_{22}^i \Delta y_{t-1} + e_t^u \end{aligned} \quad (4)$$

The relation between the structural shocks and the reduced form VAR residuals is crucial in identifying the structural shocks. This can be expressed as:

$$e_t = G_0 \varepsilon_t \quad (5)$$

Where $G_0 = A_0^{-1} B$ is a 2x2 matrix representing the contemporaneous effects of the one standard deviation shocks on the two variables.

It follows that one cannot make a distinction between the supply and demand equations unless we impose some qualifications for defining the shocks. In order to separate out the demand from the

¹ Here y and u are assumed to be I(1) and hence the VAR model uses first differences of the respective variables.

supply shocks, the BQ method defines a demand shock as the one that does not have any long run effect on the output level. If we denote G_0 as:

$$G_0 = \begin{bmatrix} g_{11}^0 & g_{12}^0 \\ g_{21}^0 & g_{22}^0 \end{bmatrix} \quad (6)$$

Then the long run restriction essentially implies that:

$$g_{12}^0 = -\frac{\sum_i c_{12}^i}{1 - \sum_i c_{22}^i} g_{22}^0 \quad (7)$$

Imposition of this restriction makes the SVAR system exactly identified and one can now identify the structural shocks ε_t^s and ε_t^d by using the information from the estimated reduced form VAR.

3. RESULTS

Looking at Figure 1, the individual responses of the markup to demand shocks show that a demand shock pushes down the markup for all countries, except in the case of Finland, in the short run, and in the case of Denmark, where the effect is negligible. On the other hand, a supply shock has a more mixed effect, increasing the markup in Belgium, France and Italy, but decreasing it in the first two-three quarters in Denmark, Sweden, and US.

[Figure 1]

To estimate the panel impact of demand and supply shocks on the evolution of markups over the short and medium-run, we follow the method proposed by Jorda (2005) which consists of estimating impulse response functions (IRFs) directly from local projections. For each period k the following equation is estimated on annual data:

$$Y_{i,t+k} - Y_{i,t} = \alpha_i^k + Time_t^k + \sum_{j=1}^l \gamma_j^k \Delta Y_{i,t-j} + \beta_k S_{i,t} + \varepsilon_{i,t}^k \quad (8)$$

with $k=1, \dots, 6$ (in years) and where Y represents our markup variable; $S_{i,t}$ represents either demand or supply shocks in country i at time t ; α_i^k are country fixed effects; $Time_t^k$ is a time trend; and β_k measures the impact of $S_{i,t}$ for each future period k . Since fixed effects are included in the regression the dynamic impact should be interpreted as compared to a baseline country-specific trend. The lag length (l) is set at 2, even if the results are extremely robust to different numbers of lags included in the specification. Equation (8) is estimated using the panel-corrected standard error (PCSE) estimator.

Impulse response functions are obtained by plotting the estimated β_k for $k=1, \dots, 6$, with confidence bands computed using the standard deviations of the estimated coefficients.²

An alternative way of estimating the dynamic impact of demand and supply shocks is to estimate an ARDL equation of changes in the markup and demand and supply shocks and to compute the IRFs from the estimated coefficients (Cerra and Saxena, 2008). However, the IRFs derived using this approach tend to be sensitive to the choice of the number of lags this making the IRFs potentially unstable. In addition, the significance of long-lasting effects with ARDL models can be simply driven by the use of one-type-of-shock models (Cai and Den Haan, 2009). This is particularly true when the dependent variable is highly persistent, as in our analysis. In contrast, the approach used here does not suffer from these problems because the coefficients associated with the lags of the change in the dependent variable enter only as control variables and are not used to derive the IRFs, and since the structure of the equation does not impose permanent effects.

Finally, confidence bands associated with the estimated IRFs are easily computed using the standard deviations of the estimated coefficients and Montecarlo simulations are not required.

[Figure 2]

In Figure 2 we can observe, and distinctly confirm, that for the panel as a whole there is a statistically significant negative impact of demand shocks on mark-ups that extends up to six quarters. On the other hand, the markup also reacts to a supply shock, although the result, in this case, is much less precisely estimated. Once again, our findings are broadly in line with previous studies. Finally, results were subjected to several robustness checks.³

4. CONCLUSION

We have assessed the effect on markups of demand and supply shocks in OECD countries using a SVAR identification. Our results show that the responses of the markups to demand shocks show are

² While the presence of a lagged dependent variable and country fixed effects may in principle bias the estimation of γ_j^k and β_k in small samples (Nickell, 1981), the length of the time dimension mitigates this concern. The finite sample bias is in order of $1/T$, where T in our sample is 38.

³ These include accounting for potential endogeneity by re-estimating the IRFs using GMM estimators. Moreover, equation (8) was re-estimated by including time fixed effects to control for specific time shocks. The results (omitted) for this specification remain statistically significant and broadly unchanged. Furthermore, a possible bias from estimating equation (8) using country-fixed effects is that the error term of the equation may have a non-zero expected value, due to the interaction of fixed effects and country-specific arrival rates of shocks. This would lead to a bias in the estimates that is a function of k . To address this issue equation (8) was re-estimated by excluding country fixed effects. The results (omitted) suggest that this bias is negligible.

negative, pushing down the markup for most countries (which we also confirmed in a panel). On the other hand, a supply shock has a more mixed effect on the markup, being also more negative in the case of the panel analysis. Such effects via demand and supply shocks are a potential explanation for the counter-cyclicality of markups found in the literature.

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Figure 1. Individual Responses of Markup to Demand and Supply Shocks

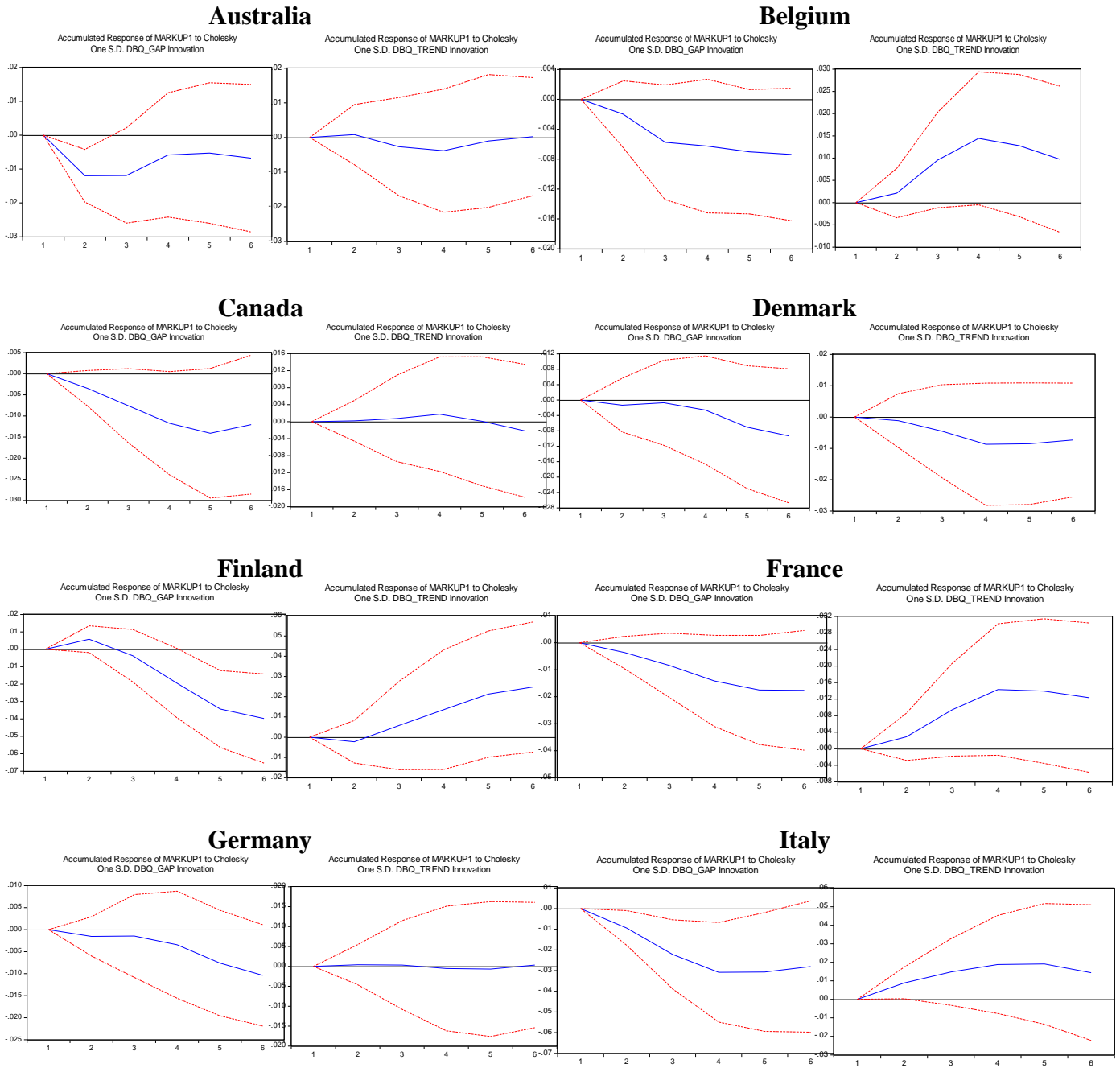
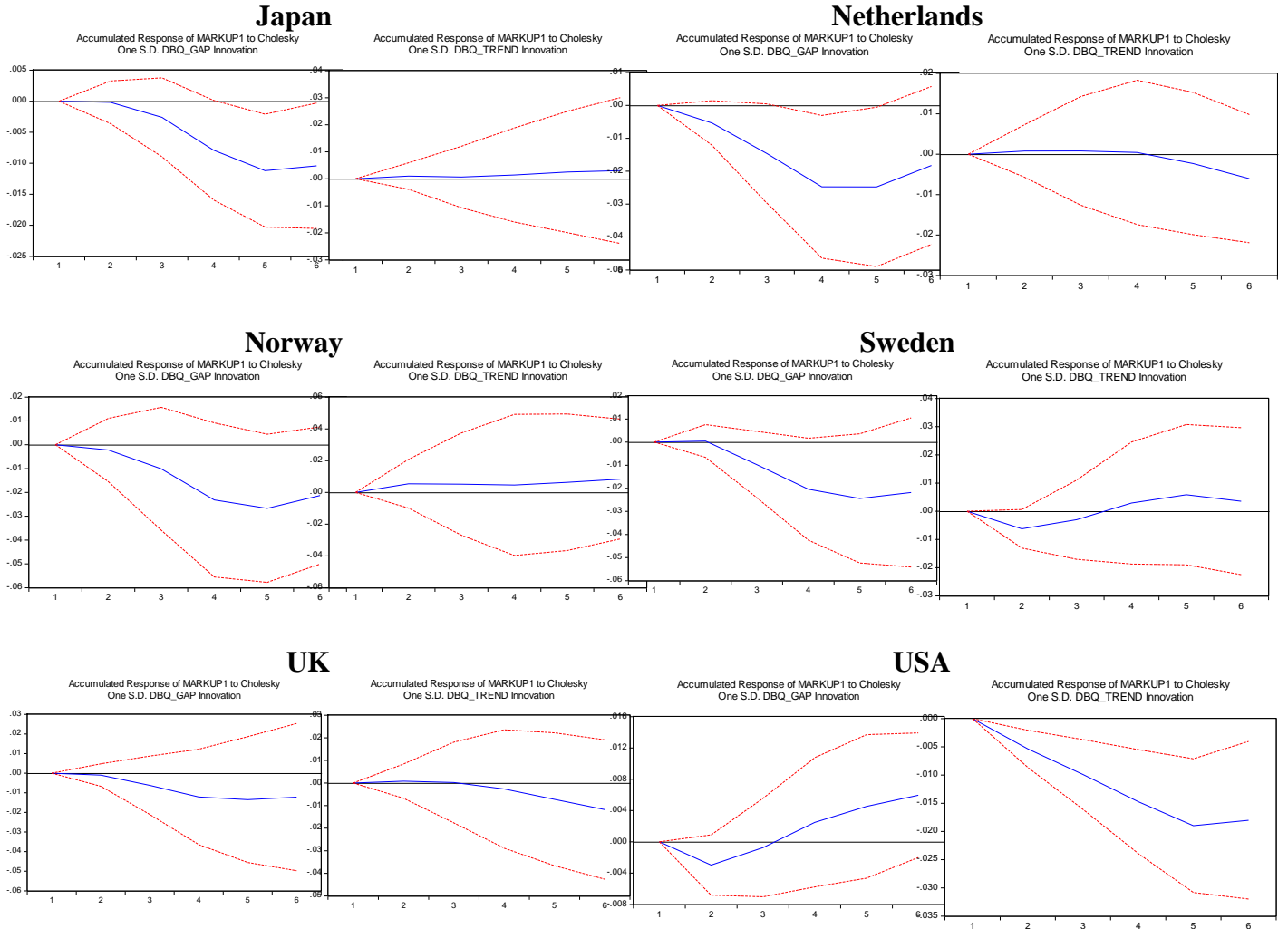
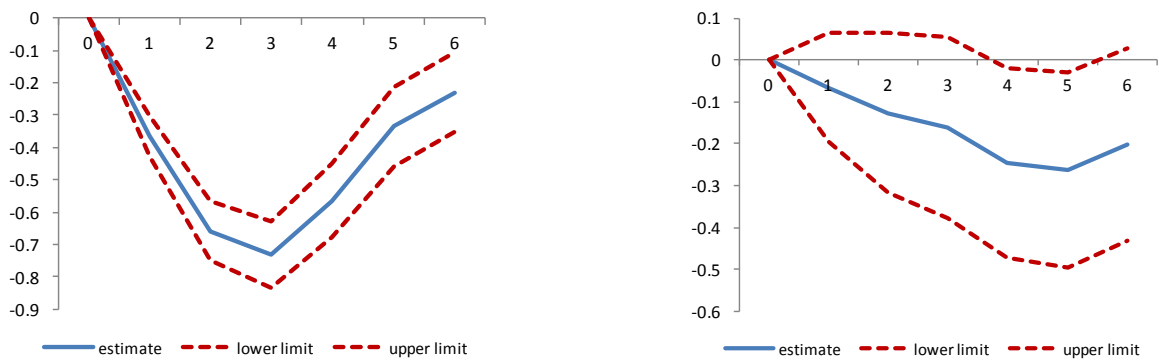


Figure 1. (cont.)



Note: Responses to a one-standard deviation shock. For each country, the left (right) char represents the impulse of mark-ups to demand (supply) shocks.

Figure 2. Panel Responses of Markup to Demand and Supply Shocks-Local Projection Estimator



Note: Dotted lines equal one standard error confidence bands. See main text for further details.