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Towards an explanation of cross-country asymmetries in monetary transmission

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Abstract

I quantify the importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission. To do so, I determine how closely the impulse responses to a monetary policy shock obtained from country-specific vectorautoregressive (VAR) models and a non-standard panel VAR model match. In the country-specific VAR models, the impulse responses vary across countries in an unrestricted fashion. In the panel VAR model, the impulse responses also vary across countries, but only to the extent that countries differ regarding their financial structure, labor market rigidities and industry mix. For a sample of 20 industrialized countries over the time period from 1995 to 2009, I find that up to 70% (50%) of the cross-country asymmetries in the responses of output (prices) to a monetary policy shock can be accounted for by cross-country differences in financial structure, labor market rigidities and industry mix. While in the short run asymmetries in the output responses arise mainly due to cross-country differences in industry mix, in the medium and long run differences in financial structure and labor market rigidities gain more importance. Moreover, cross-country differences in industry mix appear to be of rather minor importance for cross-country asymmetries in the transmission of monetary policy to prices.

Keywords: Monetary Transmission, Financial Structure, Labor Market Rigidities, Industry Mix, Panel VAR, Heterogeneity.
JEL-Classification: C33, C51, E44, E52.

Non-technical Summary

The structures of many industrialized economies have changed markedly over the last decades. Globalization nd technological progress have fostered structural change, labor market deregulation and financial market development as well as integration. These changes may have important implications for the conduct of monetary policy, as labor market rigidities may dampen the sensitivity of inflation to monetary policy and enhance that of output; output and prices may be more sensitive to monetary policy changes the larger the share of an economy's output that is accounted for by durable goods manufacturing, that is, goods with high interest rate sensitivity of demand; the intensity of competitive pressures in an economy's banking system may affect the speed and extent of pass-through of changes in the policy rate to market rates faced by households and firms; the more important external financing in an economy, the stronger the amplification of the effects of monetary policy through the credit channel. The trends in structural change, labor market deregulation and financial market development observed over the last decades, therefore, may markedly alter the transmission mechanism of monetary policy. The question is: how strongly may these developments alter the monetary transmission mechanism? How important are financial structure, labor market rigidities and industry mix for the monetary transmission mechanism? In this paper, I analyze a sample of 20 industrialized countries over the time period from 1995 to 2009 finding that 70% (50%) of the cross-country asymmetries in monetary transmission can be accounted for by cross-country differences in financial structure, labor market rigidities and industry mix. I obtain these results by comparing the impulse responses of output and prices to a monetary policy shock from two models. In the first model, the impulse responses vary in an unrestricted fashion across countries. In the second model, the impulse responses vary across countries in a restricted fashion, namely only to the extent and only if countries differ regarding their financial structure, labor market rigidities and industry mix. The better the impulse responses from model one can be replicated by the impulse responses of model two, the more the cross-country asymmetries in monetary transmission are accounted for by

cross-country differences in financial structure, labor market rigidities and industry mix. The results suggest that policies in a currency union that aim at harmonizing labor market regulations, fostering structural change and financial integration as well as innovation have a large potential to reduce asymmetries in monetary transmission. Moreover, since the driving forces of structural change, labor market deregulation and financial market development observed in the last decades are likely to persist, these results suggest that the monetary transmission mechanism is likely to change in the future. Finally, the results of this paper suggest that financial structure, labor market rigidities and industry mix should be incorporated in theoretical business cycle models which are used for policy advise.

Nicht-technische Zusammenfassung

Die ökonomischen Strukturen vieler entwickelter Volkswirtschaften haben sich in den letzten Jahrzehnten merklich verändert. Infolge der Globalisierung der Produktion und des durch den technologischen Fortschritt getriebenen industriellen Strukturwandels sind zahlreiche Reformen zur Flexibilisierung der Arbeitsmärkte durchgeführt worden. Globalisierung und technologischer Fortschritt haben auch die Weiterentwicklung der Finanzmärkte befördert. Die Veränderungen in diesen ökonomischen Strukturen ist potenziell von großer Bedeutung für die Geldpolitik. Arbeitsmarktrigiditäten können die Sensitivität der Inflation auf geldpolitische Impulse vermindern und die der Produktion verstärken. Der Anteil langlebiger Konsum- und Investitionsgüter an der gesamtwirtschaftlichen Produktion beeinflusst deren Zinssensitivität. Die Wettbewerbsintensität im Bankensystem einer Ökonomie bestimmt, wie schnell und wie stark Zinsveränderungen der Zentralbank an Haushalte und Firmen weitergegeben werden. Das Volumen an Kreditfinanzierung in einer Ökonomie beeinflusst die Verstärkung geldpolitischer Impulse durch den Kreditkanal. Die oben beschriebenen Entwicklungen in den ökonomischen Strukturen der entwickelten Volkswirtschaften verändern also prinzipiell den monetären Transmissionsmechnismus. Die Frage ist: Wie stark? Wie wichtig sind Arbeitsmarktrigiditäten, Finanzmarkt- und Industriestruktur für den monetären Transmissionsmechanismus? Dieser Frage widmet sich die vorliegende Arbeit. Die Untersuchung einer 20 Länder und den Zeitraum von 1995 bis 2009 umfassenden Stichprobe liefert das Ergebnis, dass bis zu 70% (50%) der Länderasymmetrien in der Transmission geldpolitischer Impulse auf die Produktion (das Preisniveau) auf Unterschiede in Arbeitsmarktrigiditäten, Finanzmarkt- und Industriestruktur zurückgeführt werden können. Diese Ergebnisse basieren auf dem Vergleich der Impuls-Antwort-Folgen der Produktion und des Preisniveaus aus zwei Modellen. Im ersten Modell variieren die Impuls-Antwort-Folgen in einer unrestringierten Weise über die Länder. Im zweiten Modell variieren die Impuls-Antwort-Folgen nur dann über die Länder, wenn die Länder sich hinsichtlich ihrer Arbeitsmarktrigiditäten, Finanzmarkt- und Industriestruktur unterscheiden. Umso besser die Impuls-Antwort-Folgen aus dem ersten Modell durch jene aus dem zweiten Modell repliziert werden können, desto mehr der Länderasymmetrien in der Transmission geldpolitischer Impulse auf die Produktion (das Preisniveau) können auf Unterschiede in Arbeitsmarktrigiditäten, Finanzmarkt- und Industriestruktur zurückgeführt werden. Die Ergebnisse der vorliegenden Arbeit implizieren, dass Politikmaßnahmen, die auf Arbeitsmarktharmonisierungen und die Förderung von Finanzmarktintegration sowie Strukturwandel abzielen, in Währungsgemeinschaften ein großes Potenzial haben, Länderasymmetrien in den monetären Transmissionsmechanismen substantiell zu verringern. Zudem implizieren die Resultate der vorliegenden Arbeit, dass sich der monetäre Transmissionsmechanismus in der Zukunft merklich verändern könnte, sollten sich Arbeitsmarktflexibilisierungen, Strukturwandel und Finanzmarktentwicklung fortsetzen. Schließlich bedeuten die Ergebnisse der vorliegenden Arbeit, dass Arbeitsmarktrigiditäten, Finanzmarkt- und Industriestruktur Kernelemente theoretischer Modelle zur Analyse wirtschaftlicher Schwankungen sein sollten, die für die Politikberatung verwendet werden.

Contents

Towards an Explanation of Cross-Country Asymmetries in Monetary Transmission[1]

1 Introduction

In this paper, I quantify the importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission. In a sample of 20 industrialized countries over the time period from 1995 to 2009 I find that up to 70% of the asymmetries in the responses of output and up to 50% in the responses of prices to a monetary policy shock can be accounted for by cross-country differences in financial structure, labor market rigidities and industry mix. I also find that while in the short run asymmetries in the output responses arise mainly due to cross-country differences in industry mix, in the medium and long run differences in financial structure and labor market rigidities gain more importance. Moreover, cross-country differences in industry mix appear to be of rather minor importance for cross-country asymmetries in the transmission of monetary policy to prices. These results suggest that policies aimed at harmonizing labor markets and fostering financial integration as well as structural change may markedly reduce asymmetries in monetary transmission in currency unions. Moreover, these results are important because they point to potentially large variations in the future monetary transmission mechanism arising through financial market development, labor market reforms in the face of globalization and structural change. Finally, the results suggest that financial structure, labor market rigidities and industry mix should be key elements of any theoretical business cycle model used for policy advice.

Numerous papers have attempted to identify the determinants of the monetary transmission mechanism by exploiting asymmetries in the effects of monetary policy on output and prices across countries (or regions and/or industries). In particular, the standard approach is to regress a feature of countries' impulse responses to a monetary policy shock (typically the maximum or the cumulated response) on time averages of countries'

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structural characteristics (see Carlino and DeFina, 1998; Hayo and Uhlenbrock, 1999; Mihov, 2001; Arnold and Vrugt, 2004; Dedola and Lippi, 2005; Peersman and Smets, 2005).[2] This approach is subject to several problems. First, it does not exploit the time-series variation in countries' structural characteristics to identify the determinants of the monetary transmission mechanism. This is inefficient, as at least financial structure, labor market rigidities and industry mix do display variation over time, see Section 3. The panel VAR model employed in this paper does exploit the time-series variation in countries' structural characteristics and should, therefore, pin down more precisely the importance of financial structure, labor market rigidities and industry mix for the monetary transmission mechanism. Second, the standard approach focuses only on a few of the features of the monetary transmission mechanism. However, besides the maximum and the cumulated impulse response to a monetary policy shock routinely examined in the existing literature, other important features of the monetary transmission mechanism involve the persistence of the response or the time it takes until the maximum response is reached. In the panel VAR model employed in this paper, the entire shape of the impulse responses of output and prices to a monetary policy shock is conditioned on countries' structural characteristics. Third, because the standard approach focuses on identifying the determinants of the monetary transmission mechanism rather than assessing their quantitative importance, it provides no guidance to policy in a currency union as to how large the returns of different harmonization policies (in terms of reducing asymmetries in monetary transmission) are. In contrast, the purpose of this paper is to quantify the importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission.

This paper also contributes to the literature on heterogeneity in cross-country panel data models. Panel datasets are appealing because they allow us to combine information stemming both from the cross-section and the time-series dimension. In order to exploit both dimensions, one needs to pool—that is to impose homogeneity of—at least one slope coefficient across cross-sectional units. However, pooling in dynamic panel data models when the dynamics are in fact cross-section specific entails even asymptotically a heterogeneity bias. Pesaran and Smith (1995) propose the mean-group estimator as a solution, in which the panel nature of the data set is exploited for estimation of the mean of the slope co-

[2] A different approach is pursued by Assenmacher-Wesche and Gerlach (2008), who split their sample based on the value of one structural characteristic and compare the averages of the impulse responses across country subsamples. The results of this approach may be hard to interpret, as one cannot control for more than one structural characteristic at a time. This is because the full country sample in this type of analysis is rather small (about ten to fifteen countries), so that sample splits based on more than one structural characteristic will result in country subsamples too small for averaging to produce reliable estimates.

efficients by averaging cross-section specific estimates. The mean-group approach treats heterogeneity as a nuisance, and therefore does not lend itself to an individual analysis of the underlying cross-sectional units. Unfortunately, most policies investigated in empirical cross-country studies are country specific. For example, national governments decide whether or not to embrace international trade liberalization. It is, therefore, unclear whether the mean-group approach to heterogeneity can be of any help for many of the problems national policymakers face. A different approach to heterogeneity is taken by Canova and Ciccarelli (2009), who propose to model slope heterogeneity by unobserved cross-section specific factors. The advantage of their approach is that efficiency gains stemming from the panel nature of the data set can be achieved, even if the dynamics in the data are allowed to be cross-section specific. However, the approach of Canova and Ciccarelli (2009) does not allow us to learn about the sources of heterogeneity, and it assumes heterogeneity to be random across countries. In contrast, Loayza and Raddatz (2007) introduce a stylized panel VAR model which exploits the panel nature of the data set and which links heterogeneity systematically to countries' observed structural characteristics. Similarly, Binder and Offermanns (2007) suggest a single-equation panel error-correction model in which the long-run dynamics depend on countries' observed structural characteristics. In this paper, I build on the systematic generalization of the approach of Loayza and Raddatz (2007) laid out in Georgiadis (2011) as well as the work of Binder and Offermanns (2007) to set up a panel VAR framework that allows to quantify the extent to which a given set of structural characteristics accounts for the cross-country asymmetries in monetary transmission.

The remainder of this paper is organized as follows: Section 2 presents the empirical evidence on cross-country asymmetries in the monetary transmission mechanism. In Section 3, I report empirical evidence on cross-country differences in financial structure, labor market rigidities and industry mix, discuss the mechanisms through which these structural characteristics may affect monetary transmission, and document that these structural characteristics are systematically related to cross-country asymmetries in monetary transmission. In Section 4, I motivate the design of the panel VAR model employed in this paper, lay out how impulse responses can be constructed and describe the empirical model specification. Section 5 presents results and Section 6 robustness checks. Finally, Section 7 concludes.

2 Cross-Country Asymmetries in Monetary Transmission

There has been extensive empirical work on cross-country asymmetries in monetary transmission, especially during the run-up and the first few years of the European Monetary Union (EMU); see Table 10 for an overview. If anything, the consensus in this literature has been that cross-country asymmetries in monetary transmission are likely to exist, but that the responses of output and prices to a monetary policy shock are too imprecisely estimated to make reliable statements about how large these asymmetries are. Most of the literature on cross-country asymmetries in monetary transmission has used VAR models, and I build on this framework in this paper as well. In particular, I start by estimating parsimonious, country-specific, but identical VAR models[3]

$$\boldsymbol{y}_{it} = \boldsymbol{\delta}_i + \sum_{j=1}^{p} \boldsymbol{A}_{ij} \cdot \boldsymbol{y}_{i,t-j} + \sum_{j=0}^{q} \boldsymbol{D}_{ij} \cdot \boldsymbol{x}_{i,t-j} + \boldsymbol{u}_{it}, \quad \boldsymbol{u}_{it} \overset{i.i.d.}{\sim} (\boldsymbol{0}, \boldsymbol{\Sigma}_{u,i}), \tag{1}$$

where $i = 1, 2, \ldots, N$ indexes countries, $t = 1, 2, \ldots, T$ indexes time, $\boldsymbol{y}_{it}$ is a $K \times 1$ vector of endogenous variables, $\boldsymbol{x}_{it}$ is an $M \times 1$ vector of exogenous variables, $\boldsymbol{u}_{it}$ is a vector of serially uncorrelated reduced-form disturbances and $\boldsymbol{A}_{ij}, \boldsymbol{D}_{ij}$ are $K \times K$ and $K \times M$ coefficient matrices, respectively. The vector of endogenous variables $\boldsymbol{y}_{it}$ includes the logarithm of real GDP, the logarithm of the price level and a three-month money market rate. The vector of exogenous variables $\boldsymbol{x}_{it}$ includes the Commodity Research Bureau's index of commodity prices to account for interest rate increases in anticipation of supply-side shocks. To conserve degrees of freedom (in particular because the panel VAR model laid out in Section 4 is highly parameterized), I include only six lags of the endogenous variables in the model, $p = 6$, and only the contemporaneous value, $q = 0$, of the exogenous variable.[4][5] Table 11 provides the list of countries included, the time

[3]One might argue that an accurate measurement of the effects of monetary policy shocks requires country-specific models (see, for example, Ehrmann, 2000) featuring different variables and/or identification assumptions. However, exploring different country-specific models also entails the risk that the results display asymmetries in monetary transmission across countries that are "artifacts of econometric methodology" (p. 43 Gerlach and Smets, 1995). In addition, specifying identical country VAR models is particularly relevant for this paper, as I intend to compare the results of the country VAR models to those of a panel VAR model, in which all countries are—except for fixed effects and the heterogeneity introduced by financial structure, labor market rigidities and industry mix—modeled identically. See, for example, Gerlach and Smets (1995), Kim (1999), Mihov (2001), and Assenmacher-Wesche and Gerlach (2008) who also use identical country VAR models for the analysis of the monetary transmission mechanism.

[4]I examine alternative choices for the lag orders in Section 6.

[5]Recent empirical work has emphasized the importance of common factors in output and inflation (see Canova, Ciccarelli and Ortega, 2007; Ciccarelli and Mojon, 2010). The commodity price index may be

periods covered and the variables used for each country. The data are monthly and for most countries cover the time period from 1995:1 to 2009:10.[6] The monetary policy shocks are identified by a Choleski decomposition with the interest rate ordered last, assuming that a monetary policy shock does not contemporaneously affect output and prices, but that the monetary authority has information on the current levels of output and prices.

Figure 1 combines the impulse responses of output and prices to an unsystematic, contractionary 100-basis-point increase in the short-term interest rate for the 20 countries in the sample. The top panel in Figure 1 displays the responses of output, starting from the period of impact of the monetary policy shock up to a horizon of 48 months. The bottom panel in Figure 1 depicts the corresponding responses of prices. Figures 2 and 3 provide the impulse responses for each country separately together with 90% asymptotic (bright shaded area) and bootstrap (dark shaded area) confidence bands. All impulse responses comply with the consensus view in the literature of how output and prices respond to a monetary policy shock (see Christiano, Eichanbaum and Evans, 1999): A delayed and persistent decline of prices, and a faster but only temporary drop in output. Tables 1 and 2 report the country rankings of the maximum, mean, and value of the impulse responses of output and prices after 48 months. The results in Figure 1 as well as in Tables 1 and 2 suggest that in the sample considered in this paper there are likely to be substantial cross-country asymmetries in the monetary transmission mechanism.[7] For example, while Ireland features a maximum decline in output of 1.9% relative to baseline in response to a contractionary monetary policy shock, output declines only by 0.3% relative to baseline in Poland. Moreover, while the average trough in output after a monetary policy shock across countries is 1.0 % below baseline, country-specific troughs fluctuate around that

able to pick up part of this cross-section dependence. A more explicit approach to addressing cross-section dependence is the global VAR model of Pesaran, Schuermann and Weiner (2004). I leave the integration of systematic state dependence into the global VAR model to future research. Another approach to accounting for cross-section dependence is the common correlated effects augmentation (CCEA) proposed by Pesaran (2006). While the results for the panel VAR model introduced below are hardly changed when using the CCEA, the corresponding impulse responses of the country VAR models turn out to be rather implausible (the results are available upon request).

[6]To obtain impulse responses that comply with the consensus view (see below), for Canada and the UK I also include the nominal effective exchange rate as country-specific exogenous variables. I obtain monthly real GDP figures from interpolation of quarterly figures using industrial production and the unemployment rate following the procedure suggested by Chow and Lin (1971). I cannot resort to industrial production as a measure of real activity, as in this paper I intend to show that the response of *aggregate* output to a monetary policy shock depends on the share of industrial output (in particular the share of durable goods manufacturing, see below).

[7]Due to the large uncertainty in the impulse response estimates, for most country pairs the null of identical impulse responses, $H_0 : \boldsymbol{ir}_i^{(j)} = \boldsymbol{ir}_s^{(j)}$, $i \neq s$ and $j = \{\text{output}, \text{prices}\}$, cannot be rejected. This is in line with the findings in the existing empirical literature on cross-country asymmetries in monetary transmission. However, the null that the country VAR models can be pooled into a fixed-effects panel VAR model can comfortably be rejected by a likelihood ratio test. These results are available on request.

average by about 53 %. This heterogeneity does not diminish after 48 months, when country-specific responses fluctuate by about 83 % around the average response across countries. Similarly, for prices, country-specific troughs and responses after 48 months fluctuate by about 54% and 57% around the country averages, respectively.

3 Countries' Structural Characteristics and Asymmetries in Monetary Transmission

In this section, I discuss how cross-country differences in financial structure, labor market rigidities and industry mix may give rise to cross-country asymmetries in monetary transmission. I show that these structural characteristics do indeed vary across the countries considered in this paper, and that they are systematically related to cross-country asymmetries in monetary transmission.

3.1 Financial Structure: Banking Sector Competitive Pressures and the Importance of Bank Credit

An economy's financial structure may affect the monetary transmission mechanism in numerous ways (see Barran, Coudert and Mojon, 1996; Dornbusch, Favero and Giavazzi, 1998; Assenmacher-Wesche and Gerlach, 2008). In this paper, I focus on the degree of competitive pressures in the banking sector and the importance of bank credit in the economy.[8] I first provide the rationale for how these aspects of financial structure may affect the monetary transmission mechanism, and then describe how they are measured in this paper.

Cottarelli and Kourelis (1994), Borio and Fritz (1995) as well as Mojon (2000) find cross-country differences in the (short-run) pass-through from policy to market rates. These differences may give rise to cross-country asymmetries in monetary transmission, as monetary policy should be more effective the more strongly and the faster changes in the policy

[8] Of course, many other aspects of financial structure may also affect the monetary transmission mechanism. For example, substantial down payment requirements may let households and firms hit borrowing constraints more frequently in periods of monetary tightening; the prevalence of fixed-rate instead of variable-rate mortgages may insulate households from changes in the stance of monetary policy; an unhealthy banking system with low capital ratios and high loan default probabilities may restrict lending more severely in periods of monetary tightening. Due to data limitations and the curse of dimensionality, capturing a much richer array of financial structure aspects is beyond the scope of this paper.

rate are passed through to interest rates faced by households and firms. When there are adjustment costs for changing lending rates, the absence of competitive pressures from other banks and/or alternative sources of financing results in a low interest rate elasticity of the demand for bank loans (see Klein, 1971; Hannan and Berger, 1991). This leaves room for banks not to pass through changes in policy rates to savers and borrowers. Indeed, Cottarelli and Kourelis (1994), Borio and Fritz (1995), Mojon (2000), Gropp, Kok Sørensen and Lichtenberger (2007) and van Leuvensteijn, Kok Sørensen, Bikker and van Rixtel (2008) find that stronger competitive pressures in a country's banking system are associated with more complete and faster interest rate pass-through.

Once bank interest rates have responded to changes in the monetary policy stance and interest rate channel effects have started to unfold, declines in output and prices are amplified through the components of the credit channel, that is, the balance sheet, the bank lending and the bank capital channel. In the balance sheet channel, a deterioration of the value of collateral and firms' net worth raises the cost of external finance, and thereby leads to a contraction in spending. In the bank lending channel, as long as banks are subject to reserve requirements, a tightening of monetary policy drains reserves from the banking system and leads to a reduction in banks' supply of loans to firms. In the bank capital channel, a monetary tightening reduces banks' profits as their refinancing costs (deposit rates) tend to increase relative to their earnings (loan rates), eventually leading to an erosion of bank capital. The decline in bank capital leads to a reduction of bank loans in order to meet regulatory capital requirements. Ceteris paribus, the more important bank credit in an economy the more strongly the effects of monetary policy on output and prices should be amplified by the credit channel components (for empirical evidence see Dornbusch et al., 1998; Cecchetti, 1999; Mihov, 2001).

For the measurement of competitive pressures in a country's banking sector, I rely on proxies of banking sector efficiency. In particular, I use the net interest margin (loan minus deposit rate) of the banking sector and bank costs relative to assets. For the measurement of the importance of bank credit in a country's economy, I use the amount of bank credit to the private sector relative to deposits and bank credit to the private sector by deposit money banks and other financial institutions relative to GDP.[9] All data for financial structure stem from the World Bank's Financial Structure Database (see Beck, Demirgüç-Kunt and Levine, 2009). To construct an index of financial structure, I

1. standardize the data on net interest margins, bank costs, bank credit relative to

[9]These measures do not include the amount of bank loans from foreign banks. These should, however, be of relatively minor importance given the evidence for a home bias in banking (see Carey and Nini, 2007; Vazquez and Garcia-Herrero, 2007).

deposits and private credit relative to GDP by subtracting the mean and dividing by the standard deviation,

2. reverse the net interest margin and bank cost data so that higher values reflect banking sectors with stronger competitive pressures,
3. and take the average across the four variables.

Higher values of the resulting financial structure index reflect financial structures that should exhibit a faster and more complete interest rate pass-through as well as more pronounced credit channel effects.

3.2 Labor Market Rigidities

In the baseline New Keynesian business cycle model, the less frequently nominal wages can be adjusted, the smaller the response of firms' marginal costs to a monetary policy shock (see, for example, Galí, 2008). As a result, the response of inflation is weaker and the real interest rate remains above its equilibrium for a longer period of time because of the moderate endogenous response of the central bank, which leads to a stronger response of output. However, in the New Keynesian model, labor market rigidities and their effects on the dynamics of output and inflation are not bound to nominal wage stickiness. Walsh (2005) finds that search and matching frictions may lead to a reduction in the elasticity of marginal costs with respect to output, and thereby to a dampened response of inflation as well as a stronger response of output to a monetary policy shock. Zanetti (2007) shows that unionized wage bargaining entails a muted and persistent response of wages to a monetary policy shock, implying a muted and gradual response of inflation as well as a strong and persistent response of output. Christoffel and Kuester (2008) find that inflation displays a dampened response to a monetary policy shock when there are fixed costs associated with maintaining existing job relationships. Campolmi and Faia (2010) document that inflation responds more strongly to monetary policy shocks in countries with lower replacement rates and higher employment protection. Lechthaler, Merkl and Snower (2010) demonstrate that hiring and firing costs render the responses of inflation and output to a monetary policy shock more persistent. While the precise mechanisms depend on the model specification, in general economies with more rigid labor markets should display stronger and more persistent responses of output as well as more muted, but potentially more persistent responses of inflation to a monetary policy shock.

As a proxy for the degree of a country's labor market rigidity, I use the Strictness of Employment Protection indicator provided by the Organization for Economic Development and Cooperation (OECD, see Venn, 2009). This indicator is compiled from 21 items covering three aspects of employment protection: individual dismissal of workers with regular contracts (notification and consultation requirements, notice periods and severance pay, compensation and reinstatement in case of dismissal contestation), additional provisions for collective dismissals (additional delays, costs, notification procedures) and the regulation of temporary contracts (pertaining to the operations of temporary work agencies).[10]

3.3 Industry Mix

Countries may display asymmetric responses to monetary policy shocks if their sectoral composition and sectors' sensitivity to monetary policy are different. All else equal, in countries with a large share of output accounted for by industries producing durable goods, the interest rate sensitivity of aggregate demand should be high, and output as well as prices should display strong responses to a monetary policy shock through the interest rate channel. Bernanke and Gertler (1995) find that durable consumption expenditures and residential investment drop more strongly than non-durable consumption and business fixed investment in response to a monetary policy shock. Carlino and DeFina (1998), Mihov (2001) and Arnold and Vrugt (2004) find that asymmetries in monetary transmission are partially explained by differences in the share of total output accounted for by manufacturing. Dedola and Lippi (2005) and Peersman and Smets (2005) find that within the manufacturing sector, sectors that produce durable goods feature stronger responses to monetary policy shocks. I use the share of total value added by durable goods-producing industries in the manufacturing sector to capture cross-country differences in the interest rate sensitivity of aggregate demand.[11] The data stem from the OECD's Structural Analysis Database (see OECD, 2010).

[10]Measuring the degree of labor market rigidity at the cross-country level is a non-trivial task. Numerous frictions such as wage stickiness, employment protection legislation, the power of unions and the efficiency of labor agencies may also affect the degree of labor market rigidity. Moreover, even within a country there may be inter-sectoral differences in labor market regulation. The OECD's Strictness of Employment Protection indicator does not account for most of these issues, but is, in contrast to other indices of labor market rigidities, available for a broad set of countries over a reasonably long time period. In Section 6 , in order to document that the results do not hinge on the use of the Strictness of Employment Protection indicator I consider an alternative index of labor market rigidities that covers fewer countries and a shorter time period.

[11]I adopt the classification of durable goods-producing sectors from Dedola and Lippi (2005). The sectors are wood and products of wood and cork, other non-metallic mineral products, basic metals and fabricated products, machinery and equipment as well as transport equipment.

3.4 Countries' Structural Characteristics and Asymmetries in Monetary Transmission: The Standard Approach

The left-hand side panels in Figure 4 depict the time averages of the financial structure index, the Strictness of Employment Protection indicator and the share of total value added by durable goods-producing industries in the manufacturing sector. Larger values for the financial structure index reflect financial systems in which bank credit is more important and in which banking sector competitive pressures are stronger. The Strictness of Employment Protection indicator is bounded between zero and four with larger values reflecting more rigid labor markets. Table 3 reports the maximum, minimum, standard deviation and average values of the financial structure index, the Strictness of Employment Protection indicator and the share of total value added by durable goods-producing industries in the manufacturing sector across countries. For the time period from 1995 to 2009, Ireland had the highest average value of the financial structure index, Portugal had the most rigid labor markets and Korea featured the highest share of durable goods manufacturing in total output. Overall, the left-hand side panels in Figure 4 suggest that there has been a sizeable degree of heterogeneity in financial structure, labor market rigidities and industry mix across countries. The right-hand side panels in Figure 4 depict the Hodrick-Prescott-filtered evolution over time.[12] In addition to the sizeable cross-country differences, there has also been substantial variation in financial structure, labor market rigidities and industry mix within countries over time. For example, the importance of bank credit has substantially expanded in Ireland and Denmark. Belgium and Germany have markedly removed labor market rigidities. While South Korea has become considerably more dependent on durable goods manufacturing, production has moved to other sectors in Australia and Spain. Figure 4 suggests that both the cross-sectional spread and the time-series variation in countries' structural characteristics should be useful in quantifying the importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission.

In order to confirm that financial structure, labor market rigidities and industry mix are systematically related to cross-country asymmetries in monetary transmission, I follow the standard approach (see Carlino and DeFina, 1998; Dedola and Lippi, 2005; Peersman and Smets, 2005) and regress some impulse response statistics on countries' structural

[12] All series are transformed to monthly frequency by linear interpolation before smoothing.

characteristics; that is, I estimate

$$f\left(\{\widehat{ir}_{ih}^{(j,VAR)}\}_{h=1,2,\ldots,H}\right) = a + \left(\frac{1}{T_i}\sum_{t=1}^{T_i} \boldsymbol{z}_{it}\right) \cdot \boldsymbol{b} + u_i, \qquad (2)$$

where $\widehat{ir}_{ih}^{(j,VAR)}$ denotes the estimated country VAR impulse responses of country i at horizon h, $j = \{\text{output}, \text{prices}\}$, and the scalar function $f(\cdot)$ returns the maximum contractionary response (in absolute value), the mean response, or the response at 48 months. The results for Equation (2) reported in Tables 4 and 5 suggest that financial structure and labor market rigidities are statistically significantly related to cross-country asymmetries in monetary transmission. The signs of the coefficient estimates are mostly in line with the reasoning in Sections 3.1 to 3.3.[13] For example, an increase in the financial structure index from the minimum (−1.52, see Table 3) to the maximum (1.65) value, ceteris paribus, implies a maximum output response to a monetary policy shock stronger by more than two percentage points ($-.007 \times 3.17$). Interestingly, industry mix is not statistically significant for all response statistics, and even features the wrong sign for prices. However, this might be due to the drawbacks of the standard approach discussed in the Introduction and in Section 4.

4 The Panel Conditionally Homogenous Vectorautoregressive Model

The standard approach for establishing links between cross-country differences in structural characteristics and asymmetries in the monetary transmission mechanism uses only the cross-sectional spread in countries' structural characteristics, see Section 3.4. In addition, it focuses only on a few of the features of monetary transmission (typically the maximum and the cumulated responses), and neglects other interesting features such as the persistence of monetary transmission. Moreover, existing work using the standard approach has not attempted to quantify the importance of specific sets of structural characteristics for the monetary transmission mechanism. In this paper, I address these shortcomings by using a panel VAR framework that embeds a direct link between countries' monetary transmission mechanisms and their structural characteristics. Since

[13]The results in Table 5 are not directly comparable to those of the theoretical literature on the role of labor market rigidities reviewed in Section 3.2, which pertain to inflation: A dampened but more persistent drop in inflation in response to a monetary policy shock in countries with more rigid labor markets may, but need not, be equivalent to a stronger, permanent decline in the price level.

standard panel data frameworks that account for heterogeneity typically treat it as a (random) nuisance, they cannot be used to investigate cross-country asymmetries in monetary transmission. In this section, I motivate what I will refer to as the Panel Conditionally Homogenous VAR (PCHVAR) model as a remedy to these limitations. I briefly describe how to estimate the PCHVAR model, how I construct impulse responses and lay out the empirical specification of the PCHVAR model I estimate.

4.1 The PCHVAR Model

Except for fixed effects, standard dynamic panel data models with homogenous slope coefficients do not allow for heterogeneity and are likely to be subject to heterogeneity bias. The mean-group framework proposed by Pesaran and Smith (1995) accounts for heterogeneity and provides consistent estimates of the cross-sectional means of the slope coefficients, but does not allow one to exploit the panel nature of the data set to recover country-specific dynamics. The approach of Canova and Ciccarelli (2009) models heterogeneity by unobserved factors. This approach allows one to exploit the panel nature of the data set to estimate country-specific dynamics, but does not allow one to learn about the sources of heterogeneity. In addition, the factors are assumed to be random even though parameter variation across countries is likely to be linked to country characteristics systematically. To overcome these limitations, more structure needs to be imposed on the data, and I build on the work of Loayza and Raddatz (2007) as well as Binder and Offermanns (2007) to do so. Loayza and Raddatz (2007) introduce a systematic link between cross-country heterogeneities and structural characteristics by specifying some of the autoregressive coefficients of the endogenous variables in a panel VAR model to be functions of countries' time-invariant structural characteristics. Based on the generalization laid out in Georgiadis (2011), I extend the approach of Loayza and Raddatz (2007) by relaxing the restriction that only some of the autoregressive coefficients in the panel VAR model are linked to countries' structural characteristics and by allowing the structural characteristics to vary over time.[14] Binder and Offermanns (2007) investigate a single-equation, panel error-correction model in which the short-run dynamics are fully country specific and in which the parameters of the long-run levels relationship are conditionally homogenous, namely only for countries and time periods that feature identical realizations of a single structural characteristic. I move beyond the work of Binder and Offermanns (2007) by extending the single-equation, univariate, conditional long-run homogeneity framework to

[14]Related work with different focus that also builds on the approach of Loayza and Raddatz (2007) includes Abbritti and Weber (2010), Sa, Towbin and Wieladek (2011) as well as Towbin and Weber (2011).

trivariate conditioning of short-run dynamics in a multiple-equations context.[15]

Consider a simple panel VAR model with systematic parameter variation:

$$\boldsymbol{y}_{it} \quad = \quad \boldsymbol{A}(\boldsymbol{z}_{it}) \cdot \boldsymbol{y}_{i,t-1} + \boldsymbol{u}_{it}, \qquad\qquad \boldsymbol{u}_{it} \overset{i.i.d.}{\sim} (\boldsymbol{0}, \boldsymbol{\Sigma}_u), \tag{3}$$

where $\boldsymbol{y}_{it}$ is a $K \times 1$ vector of endogenous variables, $\boldsymbol{z}_{it}$ is an $R \times 1$ vector of exogenous conditioning state variables reflecting a country's structural characteristics, and $\boldsymbol{A}(\boldsymbol{z}_{it})$ is a $K \times K$ coefficient matrix with each scalar element being a function of countries' structural characteristics $\boldsymbol{z}_{it}$. Notice that while the matrix function $\boldsymbol{A}(\cdot)$ is not country-specific, the coefficient matrix $\boldsymbol{A}(\boldsymbol{z}_{it})$ nevertheless varies across countries and over time with the realization of the country and period-specific structural characteristics. The countries' dynamics, therefore, are identical only for countries sharing the same structural characteristics; that is, the countries' dynamics are *conditionally* homogenous.

4.2 Estimation of the PCHVAR Model

The functionals collected in the coefficient matrix $\boldsymbol{A}(\cdot)$ linking countries' dynamics to their underlying structural characteristics are unknown. In order to operationalize the PCHVAR framework, I assume that each scalar coefficient functional $a_{sm}(\boldsymbol{z}_{it})$, $s, m = 1, 2, \ldots, K$, in the coefficient matrix $\boldsymbol{A}(\cdot)$ can be written as a scalar polynomial in countries' structural characteristics; that is

$$a_{sm}(\boldsymbol{z}_{it}) = \boldsymbol{\pi}(\boldsymbol{z}_{it}) \cdot \boldsymbol{\gamma}_{sm}, \tag{4}$$

where $\boldsymbol{\pi}(\boldsymbol{z}_{it}) = \left[\pi_1(\boldsymbol{z}_{it}), \pi_2(\boldsymbol{z}_{it}), \ldots, \pi_\tau(\boldsymbol{z}_{it})\right]$ is a $1 \times \tau$ vector with polynomials in $\boldsymbol{z}_{it}$, and $\boldsymbol{\gamma}_{sm} = (\gamma_{sm1}, \gamma_{sm2}, \ldots, \gamma_{sm\tau})'$ is a $\tau \times 1$ vector of polynomial coefficients. The coefficient

[15] In another paper, Binder and Offermanns (2008) set up a global VAR model in which some of the cross-country heterogeneities are linked to differences in countries' exposure to international trade and financial integration.

matrix $\boldsymbol{A}(\cdot)$ can then be written as

$$\begin{aligned}
\boldsymbol{A}(\boldsymbol{z}_{it}) &= \begin{bmatrix} \boldsymbol{\pi}(\boldsymbol{z}_{it})\cdot\boldsymbol{\gamma}_{11} & \dots & \boldsymbol{\pi}(\boldsymbol{z}_{it})\cdot\boldsymbol{\gamma}_{1K} \\ \vdots & \ddots & \vdots \\ \boldsymbol{\pi}(\boldsymbol{z}_{it})\cdot\boldsymbol{\gamma}_{K1} & \dots & \boldsymbol{\pi}(\boldsymbol{z}_{it})\cdot\boldsymbol{\gamma}_{KK} \end{bmatrix} \\
&= \begin{bmatrix} \boldsymbol{\gamma}'_{11} & \boldsymbol{\gamma}'_{12} & \dots & \boldsymbol{\gamma}'_{1K} \\ \vdots & \vdots & \ddots & \vdots \\ \boldsymbol{\gamma}'_{K1} & \boldsymbol{\gamma}'_{K2} & \dots & \boldsymbol{\gamma}'_{KK} \end{bmatrix} \cdot [\boldsymbol{I}_K \otimes \boldsymbol{\pi}'(\boldsymbol{z}_{it})] \\
&= \boldsymbol{\Gamma}\cdot[\boldsymbol{I}_K \otimes \boldsymbol{\pi}'(\boldsymbol{z}_{it})]\,. \qquad (5)
\end{aligned}$$

Using Equation (5), Equation (3) can then be written as

$$\begin{aligned}
\boldsymbol{y}_{it} &= \boldsymbol{\Gamma}\cdot[\boldsymbol{I}_K \otimes \boldsymbol{\pi}'(\boldsymbol{z}_{it})]\,\boldsymbol{y}_{i,t-1} + \boldsymbol{u}_{it} \\
&= \boldsymbol{\Gamma}\cdot\boldsymbol{\mathcal{X}}_{i,t-1}(\boldsymbol{z}_{it}) + \boldsymbol{u}_{it}, \qquad (6)
\end{aligned}$$

where $\boldsymbol{\mathcal{X}}_{i,t-1}(\boldsymbol{z}_{it})$ is a $K\tau \times 1$ vector. The model in Equation (6) is a standard multiple-equations panel data model with the same explanatory variables in every equation. Consequently, standard least-squares estimation techniques can be applied. Once the polynomial coefficients γ_{smj}, $s, m = 1, 2, ..., K$, $j = 1, 2, ..., \tau$, collected in the matrix $\boldsymbol{\Gamma}$ have been estimated, the reduced-form coefficient matrix $\boldsymbol{A}(\cdot)$ can be calculated for different values of countries' structural characteristics.[16]

4.3 Impulse Responses

Substituting recursively in Equation (3) reveals that the impulse responses in the PCHVAR model do not only depend on a country's structural characteristics in the impact period, but also on their subsequent path. Unless a country's structural characteristics do not change over time, the PCHVAR model will yield time-varying impulse responses. This complicates the comparison of the impulse responses of the country VAR model to those of the PCHVAR model, as the former features only a single, time-invariant impulse response for each country. However, if one interprets the impulse responses of the country VAR model as an average impulse response over the sample period, this problem can be addressed by constructing a corresponding statistic for the time-varying impulse responses

[16] In Georgiadis (2011), I describe the general PCHVARX(p) case with deterministic terms and exogenous variables, differences in the polynomial orders across coefficients $a_{sm}(\boldsymbol{z}_{it})$, and generalized least-squares estimation taking into account cross-sectional heteroskedasticity.

of the PCHVAR model.[17] In particular, I construct a single country-specific, average impulse response in the PCHVAR model as follows. For each country i, $i = 1, 2, \ldots, N$, and for each time period t, $t = 1, 2, \ldots, T$, I construct a country-specific and period-specific impulse response, $\{\widehat{IR}_{it}(h)\}_{h=0,1,\ldots,H}$. For this impulse response, I fix the values of country i's structural characteristics at the values observed in period t for all response horizons h, $h = 0, 1, \ldots, H$. I repeat this for all time periods t, $t = 1, 2, \ldots, T$, for country i. Then, I obtain the country-specific, average impulse response of country i as

$$\widehat{ir}_{ih}^{(PCHVAR)} \equiv T^{-1} \cdot \sum\nolimits_{t=1}^{T} \widehat{IR}_{it}(h), \tag{7}$$

which I compare to the impulse responses obtained from the country VAR models.[18]

4.4 The Estimated Model

The baseline PCHVAR model I estimate is given by

$$\boldsymbol{y}_{it} = \boldsymbol{\delta}_i + \sum_{j=1}^{p} \boldsymbol{A}_j(\boldsymbol{z}_{it}) \cdot \boldsymbol{y}_{i,t-j} + \sum_{j=0}^{q} \boldsymbol{D}_{ij} \cdot \boldsymbol{x}_{i,t-j} + \boldsymbol{u}_{it}, \quad \boldsymbol{u}_{it} \overset{i.i.d.}{\sim} (\boldsymbol{0}, \boldsymbol{\Sigma}_u), \tag{8}$$

and is analogous to the country VAR models in Equation (1). I use cubic polynomials to approximate the unknown functionals in the coefficient matrices $\boldsymbol{A}_j(\cdot)$, $j = 1, 2, \ldots, p$ without interacting the polynomials of different structural characteristics for the sake of parsimony. The same data, time periods and lag order specifications are used for estimation of the PCHVAR model as for the country VAR models.

[17] Another approach would be to estimate time-varying country VAR models using Bayesian methods (see, for example, Primiceri, 2005). The approach I pursue in this paper is conceptually more straightforward, and I leave the estimation of time-varying country-specific VAR models for future research.

[18] The results are similar for the median impulse responses within countries over time. A similar approach would use the *actual path* of the structural characteristics upon impact. However, notice that under this approach for time periods $t = T - H + 1, T - H + 2, \ldots, T$, no impulse response could be computed, as the required history would extend beyond the last time-series observation in the sample. Because it would be based on a shorter sample period, the resulting impulse response of the PCHVAR model would not be comparable to those of the country VAR models. Yet another approach would be to integrate out the path dependence of the impulse responses using Monte Carlo techniques along the lines of Koop, Pesaran and Potter (1996). I report results for this approach in Section 6.

5 Results

Figures 5 and 6 display the impulse responses of output and prices to a monetary policy shock obtained from the PCHVAR model together with 90% bootstrap confidence bands. Figures 7 and 8 combine the impulse responses obtained from the country VAR models discussed in Section 2 (solid lines) with those from the PCHVAR model (dotted lines). If financial structure, labor market rigidities and industry mix were the sole determinants of cross-country asymmetries in monetary transmission, then the impulse responses obtained from the country VAR models and those from the PCHVAR model would coincide.[19] Figure 7 suggests that for output, most of the impulse responses of the PCHVAR model match those of the country VAR models rather well. This is confirmed by Table 6 which reports the results from Wald tests of the null $H_0 : \boldsymbol{ir}_i^{(j,VAR)} = \boldsymbol{ir}_i^{(j,PCHVAR)}$, $j = \{\text{output}, \text{prices}\}$. Only for few countries can the null be rejected.[20] Moreover, Table 7 reports strong and highly statistically significant rank correlations for the maximum, the mean and the value after 48 months between the impulse responses of output obtained from the country VAR models and those from the PCHVAR model. For prices, Figure 8 seems to suggest that cross-country differences in financial structure, labor market rigidities and industry mix are less relevant for cross-country asymmetries in the responses to a monetary policy shock. Only for relatively few countries do the impulse responses of the PCHVAR model match closely those of the country VAR models. Also, despite the uncertainty in the impulse response estimates the null $H_0 : \boldsymbol{ir}_i^{(prices,VAR)} = \boldsymbol{ir}_i^{(prices,PCHVAR)}$ can be rejected for more countries than in the case of output, see Table 6. However, the rank correlations reported in Table 7 suggest that at least in the long run when prices display the strongest response, the price level impulse responses of the PCHVAR model match those of the country VAR models reasonably well.

The results presented so far show that accounting for cross-country differences in financial structure, labor market rigidities and industry mix in the PCHVAR model produces impulse responses that are similar to those obtained from country-specific VAR models. However, in order to quantify the importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission mechanism, I

[19]Of course, estimation uncertainty would have to be absent as well; see Section 6.2 for a discussion.

[20]The Wald test statistics are given by

$$W_i^{(j)} = \left(\widehat{\boldsymbol{ir}}_i^{(j,VAR)} - \widehat{\boldsymbol{ir}}_i^{(j,PCHVAR)}\right)' \cdot \left(\widehat{\boldsymbol{\mathcal{W}}}_i^{(j)}\right)^{-1} \cdot \left(\widehat{\boldsymbol{ir}}_i^{(j,VAR)} - \widehat{\boldsymbol{ir}}_i^{(j,PCHVAR)}\right), \tag{9}$$

where $j = \{\text{output}, \text{prices}\}$ and $\boldsymbol{\mathcal{W}}_i^{(j)} = Var\left(\widehat{\boldsymbol{ir}}_i^{(j,VAR)} - \widehat{\boldsymbol{ir}}_i^{(j,PCHVAR)}\right)$. I obtain both $\widehat{\boldsymbol{\mathcal{W}}}_i^{(j)}$ and the critical values for the Wald test from a bootstrap.

need to come up with a rigorous statistical measure for how closely the impulse responses obtained from the country VAR models and those from the PCHVAR model match. To do so, for a fixed response horizon h, $h = 0, 1, 2, \ldots, H$, I regress the impulse responses obtained from the country VAR models on those from the PCHVAR model

$$\widehat{ir}_{ih}^{(j,VAR)} = \alpha_h^{(j)} + \beta_h^{(j)} \cdot \widehat{ir}_{ih}^{(j,PCHVAR)} + \delta_{ih}^{(j)}, \tag{10}$$

where $j = \{\text{output}, \text{prices}\}$. The variation in the dependent variable reflects the cross-country asymmetries in monetary transmission in the data and the variation in the explanatory variable stems exclusively from cross-country differences in financial structure, labor market rigidities and industry mix. Therefore, the R-squared of the regression in Equation (10) represents the fraction of cross-country asymmetries in monetary transmission accounted for by cross-country differences in financial structure, labor market rigidities and industry mix. The top panel in Figure 9 displays the evolution of the R-squared for Equation (10) over response horizons for output. The horizontal line represents the average of the R-squareds over response horizons. The bottom panels in Figure 9 depict the evolution of the intercept and the slope estimate of Equation (10) over response horizons. The results in Figure 9 confirm the previous findings for output: Up to 70% (and on average 60%) of the cross-country asymmetries in the responses of output to a monetary policy shock can be accounted for by differences in financial structure, labor market rigidities and industry mix. Interestingly, for almost all response horizons, one cannot reject the null that the slope coefficient in Equation (10) is equal to unity and that the intercept is equal to zero. For prices, Figure 10 suggests that cross-country differences in financial structure, labor market rigidities and industry mix explain up to 50% of the cross-country asymmetries in monetary transmission at long horizons, but virtually nothing at short horizons. This is in line with the results in Table 7. The finding that the slope estimate for Equation (10) in the bottom right-hand side panel in Figure 10 is substantially above unity at long horizons while the intercept estimate is close to zero indicates that even though the PCHVAR model produces price level impulse responses that are systematically smaller in absolute value than those of the country VAR models, it still gets the relative magnitudes of the impulse responses of the country VAR models right. This is the reason why the R-squareds at long horizons in Figure 9 are high, although the price level impulse responses obtained from the country VAR models and those from the PCHVAR model do not match too closely in Figure 8. The result that cross-country asymmetries in the transmission of monetary policy to output can be explained reasonably well at short and medium horizons but not at long horizons (and vice versa for prices) is in line with most economists' beliefs about the potency of monetary policy: The classical dichotomy posits that monetary policy may affect output temporarily in the short and medium run, but not in the long run. Econometric estimates indicating a non-zero effect of monetary policy on output even in the long run are still consistent with the classical dichotomy as long as these estimates are not statistically significantly different from zero. In this case, the non-zero point estimates would be ascribed to sampling uncertainty. But if the estimates and the asymmetries in them across countries reflect sampling uncertainty only, it should not be possible to link them systematically to countries' structural characteristics, and hence the declining R-squared at long horizons in Figure 9. An analogous rationale can be put forward for prices, which due to price stickiness should respond to monetary policy only in the medium and long run.

An even better understanding of the importance of financial structure, labor market rigidities and industry mix for the monetary transmission mechanism based on the results above would be achieved if one could decompose the R-squareds in Figures 9 and 10 into the contributions of the three structural characteristics. It might appear as if this could be carried out easily by considering financial structure, labor market rigidities and industry mix individually one at a time as scalar conditioning state variable z_{it} in Equation (8). Unfortunately, as in any linear regression model, dropping two of the three structural characteristics will in general not result in an exact decomposition of the R-squareds in Figures 9 and 10. In addition, because the impulse responses in the PCHVAR model are multiplicative rather than additive functions of the structural characteristics, the sum of the R-squareds from the univariate PCHVAR models may even fall short of the R-squared from the baseline model in which the structural characteristics are considered jointly. However, considering financial structure, labor market rigidities and industry mix one at a time as conditioning state variable in Equation (8) may still be helpful to get an idea about the relative importance of the three structural characteristics. Moreover, it may point to differences in the response horizons at which the structural characteristics affect the monetary transmission mechanism. Figures 11 and 12 present the evolution of the R-squared for Equation (10) over response horizons when financial structure, labor market rigidities and industry mix are considered one at a time as univariate conditioning state variable in Equation (8). For output, two observations stand out. First, at medium to long horizons, only financial structure and labor market rigidities appear to be of (similar) importance for cross-country asymmetries in monetary transmission. Second, in the short run, industry mix appears to be much more important than financial structure and labor market rigidities. This is in line with the view that the monetary transmission mechanism first works through the interest rate channel and then gets amplified through credit channel effects and other frictions such as labor market rigidities. This result could not have been obtained from the standard approach; see the statistically insignificant coefficients for industry mix in Table 4. For prices, it appears that only financial structure and labor market rigidities appear to be of (similar) importance for cross-country asymmetries in monetary transmission in the medium and long run.

Another insightful exercise is to examine whether financial structure, labor market rigidities and industry mix are of different importance for cross-country asymmetries in monetary transmission across country groups. Figure 13 presents the evolution of the R-squared for Equation (10) for the full sample (solid lines), Continental European (dashed lines) and Anglo-Saxon countries (dash-dotted lines). The results in Figure 13 suggest that financial structure, labor market rigidities and industry mix account for a much larger fraction of the cross-country asymmetries in the transmission of monetary policy to prices across the Anglo-Saxon countries than across the Continental European countries. For the asymmetries in the transmission of monetary policy to output, it appears that financial structure, labor market rigidities and industry mix are of similar importance both for the Continental European and Anglo-Saxon countries at medium horizons, but significantly more important for Anglo-Saxon countries at short horizons. However, at least the results fort the Anglo-Saxon countries should be taken with caution due to the small sample size.

6 Robustness and Discussion

In this section, I present results for alternative specifications of the country VAR and the PCHVAR models, for an alternative way of constructing impulse responses for the PCHVAR model, and discuss several more general issues regarding the empirical approach taken in this paper.

6.1 Robustness

The parsimonious three-variable country VAR model does not account for the exchange rate channel that may be important for some of the more open countries considered in this paper. Moreover, the baseline country VAR models do not account for the fact that central banks may monitor monetary aggregates when setting policy rates. Finally, while in most of the countries considered in this paper monetary policy is implemented by targeting the overnight money market rate (see Table 9 and the discussion below), in the baseline specification I use the three-month money market rate to reflect monetary policy. I do so in order to ensure the consistency of the data used: the overnight money market rate series from the OECD's Main Economic Indicators for the euro area countries are available only from 1999 and those from the IMF's International Financial Statistics end in 1998 for most euro area countries; for the few countries and time periods for which both the OECD and the IMF provide data for the euro area countries, these do not exactly match. Figures 14 and 15 display the impulse responses of output and prices to a monetary policy shock obtained from country VAR models analogous to those in Equation (1), but including the nominal effective exchange rate (dash-dotted lines) or M3 (dashed lines) as additional endogenous variables. Figures 14 and 15 also provide the impulse responses from country VAR models analogous to those in Equation (1) but with an overnight rate (dotted lines) instead of a three-month money market rate series obtained from a combination of data from the OECD and the IMF used to reflect monetary policy.[21] The solid lines represent the baseline country VAR model impulse responses from Figure 1. The results of the alternative specifications in Figures 14 and 15 are similar to the baseline results.

In the baseline country VAR model specifications, the lag orders of the endogenous variables are set to six and those of the exogenous variables to zero. The reason for not determining the lag orders optimally according to some information criterion is that I compare the results of the country VAR models to those of the PCHVAR model, in which all countries feature the same lag order by construction. In Figures 16 and 17, the impulse responses of the country VAR models estimated with a lag order of nine for the endogenous variables (dash-dotted line) and three for the exogenous variables (dashed line) are shown. The solid lines represent the baseline impulse responses of the country VAR models. The results in Figures 16 and 17 suggest that the impulse responses of output and prices to a monetary policy shock in the country VAR models are similar across alternative lag order specifications.

For the euro area countries, the inclusion of euro area aggregate output and prices may be necessary for the identification of monetary policy shocks. Figures 18 and 19 display the impulse responses of output and prices to a monetary policy shock obtained from country VAR models analogous to those in Equation (1), but with euro area aggregate output and prices included as additional endogenous variables (dash-dotted lines).[22] The solid lines represent the baseline country VAR impulse responses. Except for the responses of Spain

[21] The exchange rate is ordered last, as it should contemporaneously respond to monetary policy shocks; money is ordered before the interest rate. There are no results for the specification with money for the euro area countries because national M3 data spanning from 1995 to 2009 are not available. I merge the overnight money market rate series for the euro area from 1999 to 2009 from the OECD's Main Economic Indicators with the overnight call money rate series from the IMF's International Financial Statistics for the time period from 1995 to 1998.

[22] Euro area output and prices are ordered first and second. Notice that since the harmonized index of consumer prices is available only from 1996 onwards, for Figures 18 and 19 I re-estimate the baseline model for the time period from 1996 to 2009. Moreover, in order to reduce the parametrization once the number of endogenous variables is increased to five instead of three as in the baseline specification, I reduce the lag order of the endogenous variables to three.

(with rather implausible impulse responses) and the price level responses of Denmark, Ireland and Portugal, the results with euro area aggregates are again similar to those from the baseline specification.

It is plausible to assume that financial structure, labor market rigidities and industry mix evolve rather slowly over time and are to a large extent determined by technological as well as political economy factors. For example, La Porta, Lopez-de-Silanes, Shleifer and Vishny (1997) argue that differences in the importance of bank credit can be traced back to whether a country's legal system has British, French, German or Scandinavian origin. Nevertheless, it is important to ensure that the results in this paper are not driven by disregarding potential feedback between monetary policy and economic activity on the one hand and countries' structural characteristics on the other hand. Figures 20 and 21 display impulse responses of output and prices to a monetary policy shock obtained from the baseline country VAR models in Equation (1) augmented by financial structure, labor market rigidities and industry mix as additional endogenous variables (one at a time to conserve degrees of freedom).[23] The solid lines represent the baseline country VAR model impulse responses from Figure 1. Except for the output response of Austria when including financial structure or the share of durable goods manufacturing in total output as well as Ireland when including the share of durable goods manufacturing in total output, the results of the alternative specifications in Figures 20 and 21 are similar to the baseline results.

Table 8 reports the correlations between the impulse responses of the alternative specifications described above and the baseline results. In particular, the table reports four correlations: the correlation between the maximum responses, the responses after 48 months, the mean responses and the full responses over all horizons. All correlations are rather high. To sum up, the impulse responses obtained from the country VAR models are mostly unchanged when the lag orders are increased, the overnight money market rate is used to reflect monetary policy, or when the exchange rate, a monetary aggregate, euro area aggregate output and prices or the structural characteristics are included as additional endogenous variables.

Regarding the PCHVAR model, I examine three robustness checks. First, Figure 22 displays the evolution over response horizons of the R-squared of Equation (10) for the baseline impulse responses implied by the PCHVAR model and constructed as described in Section 4.3 (solid lines) together with those obtained from integrating out the time-series variation of the structural characteristics along the lines of Koop et al. (1996) and

[23]The structural characteristics are ordered first.

described in more detail in Appendix B (dash-dotted lines). Second, the dotted lines represent the evolution over response horizons of the R-squared of Equation (10) when the OECD's Strictness of Employment Protection indicator is replaced by the average over (standardized data on) a country's benefit replacement rate, wage bargaining co-ordination, labor tax rate, product market regulations, social benefit spending relative to GDP, union density and employment protection legislation.[24] The results are mostly unchanged when constructing the impulse responses implied by the PCHVAR model by means of Monte Carlo integration and when using an alternative, more comprehensive index of labor market rigidities. Finally, since the construction sector should also feature a relatively strong interest rate sensitivity of demand for its output, the dashed lines in Figure 22 display the evolution over response horizons of the R-squared of Equation (10) for the baseline model with the share of value added by the construction sector added to that of durable goods manufacturing as a measure of industry mix. The inclusion of the share of value added by the construction sector in the measure for industry mix appears to lower the joint explanatory power of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission, in particular at medium and long horizons. This suggests that cross-country differences in the share of value added by the construction sector are not an important determinant of cross-country asymmetries in monetary transmission.

6.2 Discussion

The main objective in this paper is to determine how closely the impulse responses obtained from country-specific VAR models match with those from the PCHVAR model. Because sampling uncertainty could lead to a discrepancy of these impulse responses even if financial structure, labor market rigidities and industry mix were the only determinants of cross-country asymmetries in monetary transmission, ideally the regression in Equation (10) should involve the *probability limits* of the impulse responses obtained from the country VAR models and those from the PCHVAR model. The probability limits are, of course, not available and I resort to finite sample estimates instead. It is, therefore, important to ensure that the findings of this paper do not arise because of the use of uncertain estimates in place of probability limits. In fact, the difference between the probability limits and the finite sample estimates enters the regression error in Equation (10). This introduces measurement error in both the dependent and the explanatory vari-

[24] I obtain the data from Berger and Heylen (2011) who gather them from various sources as well as earlier papers and update them until 2007. Since these labor market data are available only for 15 of the 20 countries in my sample, I extrapolate them until 2009 in order not to lose too many observations.

able, which can be shown to bias downward the R-squared of Equation (10); see Section A.2.[25] As a result, even if financial structure, labor market rigidities and industry mix are the only determinants of cross-country asymmetries in monetary transmission, estimation uncertainty will drive the R-squared of Equation (10) *below* unity. This can be illustrated by running a weighted least squares regression of Equation (10), in which the weights are negatively related to the magnitude of the standard errors of the corresponding impulse response estimates. Figure 23 displays the evolution of the R-squared in the baseline results (solid line) and the corresponding results for the weighted least squares regression (dashed lines). As can be seen, granting more weight to those impulse response observations that are estimated less imprecisely leads to an increase in the R-squareds.[26] The presence of estimation uncertainty in the impulse response estimates of output and prices to a monetary policy shock should, therefore, not lead to an overestimation of the quantitative importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in monetary transmission.

Several papers in the literature on cross-country asymmetries in monetary transmission (see Table 10) argue that the true impulse responses of output and prices are in fact identical across countries, and that the asymmetries displayed in Figure 1 are random and due to sampling uncertainty. This might raise the concern that the results of this paper are spurious, since the differences across countries in the impulse response estimates are the major sources of information for the identification of the importance of financial structure, labor market rigidities and industry mix for cross-country asymmetries in the monetary transmission mechanism. However, it is unlikely that the results in this paper are biased because I exploit spurious variation in impulse response estimates across countries. Think of a standard linear regression framework, and suppose that (i) there is no variation in the dependent variable, and that (ii) only an imprecise estimate of the dependent variable is available. Only if the measurement error in the dependent variable is systematically related to the explanatory variables will the latter feature statistically significant coefficients. It appears hard to make the case for cross-country differences in financial structure, labor market rigidities and industry mix being systematically related

[25] Moreover, the measurement error also biases the slope estimate towards zero and the intercept estimate up; see Equation (A.5). For output, both effects render it easier to *reject* the null that the slope is equal to unity and the intercept is equal to zero. Moreover, in the presence of measurement error, the confidence bands depicted in the bottom panels of Figure 9 are spuriously tight, which should lead to over-rejection of the null hypotheses that the intercept is equal to zero and the slope is equal to one in Equation (10); see Equation (A.9).

[26] Notice, however, that this result has to be taken with caution, as the R-squared of the weighted least squares regression is not bounded between zero and one and also depends on the precise weighting scheme.

to random variation in the estimates of countries' monetary transmission mechanisms.[27]

Another concern might be the assumption of homoskedasticity in the baseline PCHVAR model in Equation (8). While erroneously assuming homoskedastic variances does not affect consistency of estimation, in general it is restrictive for impulse responses because it implies that contemporaneous correlations of structural shocks are identical across countries. Notice, however, that when identifying monetary policy shocks using the Choleski decomposition, ordering the interest rate last implies that country-specific impulse responses to a monetary policy shock do *not* depend on whether contemporaneous correlations between shocks are allowed to differ across countries. The intuition is that when neither output nor prices are assumed to respond contemporaneously to the monetary policy shock, the contemporaneous correlation between the monetary policy shock and the shocks in the output and price level equations is zero by construction (of course, heteroskedasticity may still be relevant for the responses to the output and price level shocks). Technically, the responses of the endogenous variables to the monetary policy shock when the interest rate is ordered last are governed by the last column of the Choleski decomposition $\boldsymbol{P}_i$ of $\boldsymbol{\Sigma}_{u,i} = \boldsymbol{P}_i \cdot \boldsymbol{\Sigma}_\epsilon \cdot \boldsymbol{P}'_i$, which is invariant to heteroskedasticity in the variance matrices $\boldsymbol{\Sigma}_{u,i}$.

A last issue is that differences in central banks' institutional frameworks might invalidate imposing the same empirical model framework on all countries. Table 9 provides an overview of the monetary policy strategies of the countries considered in this paper. All countries pursue an inflation-targeting approach, either explicitly institutionalized ("full-fledged") or implicitly in connection with a price stability anchor ("eclectic"). All central banks target the overnight money market rate, with the exception of the Swiss and the Hungarian central banks, which target the three-month money market rate. Almost all central banks aim to steer money market rates by maintaining an interest rate corridor with standing lending and/or deposit facilities. Taken together, it appears that the operating frameworks of the central banks for the countries considered in this paper are sufficiently similar to justify an approach that imposes the same empirical framework on all countries.

[27]Similar reasoning applies to several other qualifications of the use of the VAR approach to cross-country asymmetries in monetary transmission in general: It is unclear whether the restriction to only output, prices and interest rates can plausibly capture the interaction of the central bank with the real economy; whether identical country-specific VAR models can produce reliable estimates of the monetary transmission if central banks' operating procedures, exchange rate regimes and other country characteristics are different; whether structural breaks due to the use of pre-EMU and post-EMU data bias the estimates; whether investigating the responses of output and prices to identical and/or country-specific shocks to monetary policy is subject to the Lucas critique. These issues are legitimate concerns, but their effect should be to render it harder to find empirical evidence for cross-country differences in financial structure, labor market rigidities and industry mix to affect monetary transmission.

7 Conclusion

In this paper, I make use of the PCHVAR model to analyze the importance of financial structure, labor market rigidities and industry mix for the monetary transmission mechanism. In the PCHVAR model, parameter heterogeneity is not viewed as a random nuisance. Instead, heterogeneity represents a source of variation that can be exploited to learn about the state dependence of, for example, policies. Moreover, in contrast to the existing empirical literature on cross-country asymmetries in monetary transmission focusing mostly on the standard approach, the PCHVAR model allows to (i) exploit the time-series variation in countries' structural characteristics to learn about their role for asymmetries in monetary transmission, (ii) to take into account the entire shape of the responses of output and prices to a monetary policy shock rather than only the maximum response, and (iii) to quantify the importance of a set of countries' structural characteristics for cross-country asymmetries in monetary transmission. I find that up to 70% (50%) of the asymmetries in the responses of output (prices) to a monetary policy shock across countries can be accounted for by jointly incorporating cross-country differences in financial structure, labor market rigidities and industry mix. A tentative decomposition of these figures into the contributions of each structural characteristic shows that while in the short run asymmetries in the output responses arise mainly due to cross-country differences in industry mix, in the medium run differences in financial structure and labor market rigidities gain more importance. Moreover, cross-country differences in industry mix appear to be of rather minor importance for cross-country asymmetries in the transmission of monetary policy to prices. These results suggest that policies aimed at harmonizing labor markets and fostering financial integration as well as structural change may markedly reduce asymmetries in monetary transmission in currency unions. Moreover, these results point to potentially large variations in the future monetary transmission mechanism arising through financial market development, labor market reforms and structural change. Finally, the results suggest that financial structure, labor market rigidities and industry mix should be key elements of any theoretical business cycle model used for policy advice.

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A The Effects of Measurement Error

In this section, I first describe the effects of measurement error in the explanatory variable on the coefficient estimates. I then proceed to a description of the effects of measurement error in the dependent and/or the explanatory variables on the regression R-squared.

A.1 The Effect of Measurement Error on the Coefficient Estimates

Consider a standard cross-section regression

$$\begin{aligned} y_i^* &= \alpha + \beta \cdot x_i^* + u_i \\ &= \boldsymbol{Z}_i^* \cdot \boldsymbol{\gamma} + u_i, \end{aligned} \tag{A.1}$$

for observations $i = 1, 2, \ldots, N$, $\boldsymbol{Z}_i^* \equiv (1, x_i^*)$, and $\boldsymbol{\gamma} \equiv (\alpha, \beta)'$. Suppose we want to estimate Equation (A.1), but we only have an inaccurate measure of the explanatory variable given by

$$x_i = x_i^* + e_i, \quad Var(e_i) = \sigma_e^2, \;\; E(e_i) = E(e_i x_i^*) = E(e_i u_i) = 0. \tag{A.2}$$

Instead of Equation (A.1), the estimated equation is thus

$$\begin{aligned} y_i^* &= \alpha + \beta \cdot x_i + u_i - \beta \cdot e_i \\ &= \alpha + \beta \cdot x_i + \delta_i \\ &= \boldsymbol{Z}_i \cdot \boldsymbol{\gamma} + \delta_i, \end{aligned} \tag{A.3}$$

with $Var(\delta_i) \equiv \sigma_\delta^2 = \beta^2 \sigma_e^2 + \sigma_u^2$. The ordinary least squares (OLS) estimator

$$\widehat{\boldsymbol{\gamma}} = \left(\sum_{i=1}^{N} \boldsymbol{Z}_i' \boldsymbol{Z}_i \right)^{-1} \left(\sum_{i=1}^{N} \boldsymbol{Z}_i' y_i^* \right), \tag{A.4}$$

entails an attenuation bias due to the non-zero correlation between the composite error, δ_i, and the regressor, x_i, $E(x_i \delta_i) = -\beta \sigma_e^2$. It can be shown that

$$plim\ \widehat{\boldsymbol{\gamma}} = \begin{bmatrix} 1 & r_x \cdot E(x_i) \\ 0 & 1 - r_x \end{bmatrix} \cdot \boldsymbol{\gamma} = \boldsymbol{A} \boldsymbol{\gamma}, \quad r_x \equiv \frac{\sigma_e^2}{\sigma_e^2 + \sigma_{x^*}^2}. \tag{A.5}$$

Given the signal-to-noise ratio, $STN \equiv \sigma^2_{x^*}/\sigma^2_e$, one can obtain a corrected coefficient estimate and variance-covariance matrix

$$\widetilde{\boldsymbol{\gamma}} = \boldsymbol{A}^{-1}\widehat{\boldsymbol{\gamma}}, \tag{A.6}$$

$$Var(\widetilde{\boldsymbol{\gamma}}) = \boldsymbol{A}^{-1}Var(\widehat{\boldsymbol{\gamma}})\boldsymbol{A}^{-1\prime}, \tag{A.7}$$

for which $plim\,\widetilde{\gamma} = \gamma$. It turns out that

$$\widetilde{\beta} = \frac{1}{1-r_x}\cdot\widehat{\beta} = \left(1+\frac{1}{STN}\right)\cdot\widehat{\beta}, \tag{A.8}$$

$$std(\widetilde{\beta}) = \frac{1}{1-r_x}\cdot std\left(\widehat{\beta}\right) = \left(1+\frac{1}{STN}\right)\cdot std\left(\widehat{\beta}\right). \tag{A.9}$$

The measurement error setting is similar but different to a generated regressors framework in the spirit of Pagan (1984):

$$y_i^* = x_i^*\cdot\delta + u_i, \tag{A.10}$$

$$x_i = x_i^* + e_i \ = q_i\cdot\alpha + e_i, \tag{A.11}$$

where x_i^* is an unobserved variable related to the observed, predetermined variables q_i. For estimation of δ, x_i^* is replaced by $\widehat{x}_i^* = q_i\widehat{\alpha}$ in the first equation. In the context of this paper, y_i^* represents the probability limit of the country VAR model impulse responses, $ir_i^{(VAR)}$, x_i^* the probability limit of the PCHVAR model impulse responses,$ir_i^{(PCHVAR)}$, and x_i the estimated PCHVAR model impulse responses, $\widehat{ir}_i^{(PCHVAR)}$. In the generated regressors framework, the coefficient estimates remain consistent, but the standard errors - unless corrected - do not take into account sampling uncertainty from the first stage regression. The difference between the generated regressors framework and the present setting is that the finite sample PCHVAR model impulse response estimate x_i is not an estimate $\widehat{x}_i^*$ of x_i^*, but rather a noisy measure.

A.2 The Effect of Measurement Error on the Goodness of Fit

Upon OLS estimation of the cross-section regression without measurement error in Equation (A.1) for the R-squared it holds that

$$R^{*2} \equiv 1 - \frac{\widehat{\sigma}^2_u}{\widehat{\sigma}^2_{y^*}} \xrightarrow{p} 1 - \frac{\sigma^2_u}{\sigma^2_{y^*}}. \tag{A.12}$$

As I show below, when the dependent and/or the explanatory variables are measured with error, the probability limits of the R-squared will be different from - namely lower than - the probability limit in Equation (A.12). Thus, even if the true fit is perfect, asymptotically the R-squared will be below unity solely due to the presence of measurement error. If one has an estimate of the signal-to-noise ratio in the measures of the dependent and the explanatory variables, at least asymptotically one can correct for the downward bias in the actual R-squared: The probability limit of the true R-squared is equal to the probability limit of the actual R-squared multiplied by one plus the inverse of the relevant signal-to-noise ratio. I first discuss the case when measurement error is present only in the dependent variable, followed by the case when only the explanatory variable is measured with error. Finally, I put pieces together and describe the effect of measurement error in both the dependent and explanatory variables on the R-squared.

A.2.1 Measurement Error in y_i^*

Suppose we want to estimate Equation (A.1) but we only have an inaccurate measure of the dependent variable, that is

$$y_i = y_i^* + w_i, \quad Var(w_i) = \sigma_w^2, \;\; E(w_i) = E(w_i y_i^*) = E(w_i u_i) = 0, \tag{A.13}$$

with $Var(y_i) \equiv \sigma_y^2 = \sigma_{y^*}^2 + \sigma_w^2$. Instead of Equation (A.1), the estimated equation is given by

$$\begin{aligned} y_i &= \alpha + \beta \cdot x_i^* + u_i + w_i \\ &= \alpha + \beta \cdot x_i^* + \delta_i, \end{aligned} \tag{A.14}$$

with $Var(\delta_i) \equiv \sigma_\delta^2 = \sigma_w^2 + \sigma_u^2$. For the R-squared of the model in Equation (A.14) it holds that

$$R_y^2 \equiv 1 - \frac{\widehat{\sigma}_\delta^2}{\widehat{\sigma}_y^2} \xrightarrow{p} 1 - \frac{\sigma_\delta^2}{\sigma_y^2}, \tag{A.15}$$

as the OLS estimator is consistent under measurement error in the dependent variable. Asymptotically, the ratio of the R-squareds of the model without any measurement error in Equation (A.1) and the model with measurement error in the dependent variable in

Equation (A.14) is given by

$$\begin{aligned} plim\ \frac{R^{*2}}{R_y^2} &= \frac{(\sigma_{y^*}^2 - \sigma_u^2)/\sigma_{y^*}^2}{(\sigma_y^2 - \sigma_\delta^2)/\sigma_y^2} \\ &= \frac{(\sigma_{y^*}^2 - \sigma_u^2)/\sigma_{y^*}^2}{(\sigma_{y^*}^2 + \sigma_w^2 - \sigma_u^2 - \sigma_w^2)/\sigma_y^2} \\ &= \frac{\sigma_y^2}{\sigma_{y^*}^2} \\ &= 1 + \frac{\sigma_w^2}{\sigma_{y^*}^2}. \end{aligned} \tag{A.16}$$

Thus,

$$plim\ R^{*2} = \left(1 + \frac{\sigma_w^2}{\sigma_{y^*}^2}\right) \cdot plim\ R_y^2. \tag{A.17}$$

A.2.2 Measurement Error in x_i^*

Suppose we want to estimate Equation (A.1) but we only have an inaccurate measure of the explanatory variable, that is

$$x_i = x_i^* + e_i, \quad Var(e_i) = \sigma_e^2, \ \ E(e_i) = E(e_i x_i^*) = E(e_i u_i) = 0, \tag{A.18}$$

with $Var(x_i) \equiv \sigma_x^2 = \sigma_{x^*}^2 + \sigma_e^2$. Instead of Equation (A.1), the estimated equation is given by

$$\begin{aligned} y_i^* &= \alpha + \beta \cdot x_i + u_i - \beta \cdot e_i \\ &= \alpha + \beta \cdot x_i + \delta_i, \end{aligned} \tag{A.19}$$

with $Var(\delta_i) \equiv \sigma_\delta^2 = \beta^2 \sigma_e^2 + \sigma_u^2$. The R-squared for the model with measurement error in the explanatory variable in Equation (A.19) is given by

$$R_x^2 \equiv 1 - \frac{\widehat{\sigma}_\delta^2}{\widehat{\sigma}_{y^*}^2}. \tag{A.20}$$

Due to the non-zero correlation between the composite error $\delta_i = u_i - \beta \cdot e_i$ and the regressor x_i, $E(x_i \delta_i) = -\beta \cdot \sigma_e^2 \neq 0$, under measurement error in the explanatory variable the OLS estimates (and therefore residuals and estimated variances) are inconsistent. Consequently, in order to find the probability limit of the R-squared of the model with measurement error in the explanatory variable given in Equation (A.19), R_x^2, one needs to determine the probability limit of $\widehat{\sigma}_\delta^2$. Using the result in Equation (A.5) in Appendix A.1, we have that

$$\begin{aligned}
plim\ \widehat{\sigma}_\delta^2 &= plim\ \frac{1}{N}\sum_{i=1}^{N}(y_i^* - \boldsymbol{Z}_i \cdot \widehat{\boldsymbol{\gamma}})^2 \\
&= plim\ \frac{1}{N}\sum_{i=1}^{N}[y_i^* - \boldsymbol{Z}_i \cdot \boldsymbol{\gamma} - \boldsymbol{Z}_i \cdot (\widehat{\boldsymbol{\gamma}} - \boldsymbol{\gamma})]^2 \\
&= plim\ \frac{1}{N}\sum_{i=1}^{N}[\delta_i - \boldsymbol{Z}_i \cdot (\widehat{\boldsymbol{\gamma}} - \boldsymbol{\gamma})]^2 \\
&= plim\ \frac{1}{N}\sum_{i=1}^{N}\delta_i^2 - 2 \cdot plim\ \frac{1}{N}\sum_{i=1}^{N}[\delta_i \boldsymbol{Z}_i \cdot (\widehat{\boldsymbol{\gamma}} - \boldsymbol{\gamma})] + plim\ \frac{1}{N}\sum_{i=1}^{N}[\boldsymbol{Z}_i \cdot (\widehat{\boldsymbol{\gamma}} - \boldsymbol{\gamma})]^2 \\
&= \sigma_\delta^2 - 2 \cdot plim\ \frac{1}{N}\sum_{i=1}^{N}\left[\delta_i \cdot (\widehat{\alpha} - \alpha) + \delta_i x_i \cdot (\widehat{\beta} - \beta)\right] \\
&\quad + plim\ \frac{1}{N}\sum_{i=1}^{N}\left[(\widehat{\alpha} - \alpha) + x_i \cdot (\widehat{\beta} - \beta)\right]^2 \\
&= \sigma_\delta^2 - 2 \cdot plim\ (\widehat{\alpha} - \alpha) \cdot plim\ \frac{1}{N}\sum_{i=1}^{N}\delta_i - 2 \cdot plim\ (\widehat{\beta} - \beta) \cdot plim\ \frac{1}{N}\sum_{i=1}^{N}\delta_i x_i \\
&\quad + plim \frac{1}{N}\sum_{i=1}^{N}\left[(\widehat{\alpha} - \alpha)^2 + 2x_i \cdot (\widehat{\alpha} - \alpha)(\widehat{\beta} - \beta) + x_i^2 \cdot (\widehat{\beta} - \beta)^2\right] \\
&= \sigma_\delta^2 - 2 \cdot r_x E(x_i)\beta \cdot E(\delta_i) - 2 \cdot (-r_x\beta) \cdot (-\beta\sigma_e^2) + plim\ (\widehat{\alpha} - \alpha)^2 \\
&= \sigma_\delta^2 - 2r_x\beta^2\sigma_e^2 + r_x^2 E(x_i)^2\beta^2 - 2r_x^2\beta^2 E(x_i)^2 + r_x^2\beta^2 E(x_i^2) \\
&= \sigma_\delta^2 - 2r_x\beta^2\sigma_e^2 - r_x^2 E(x_i)^2\beta^2 + r_x^2\beta^2 E(x_i^2) \\
&= \sigma_\delta^2 - 2r_x\beta^2\sigma_e^2 + r_x^2\beta^2\sigma_x^2 \\
&= \sigma_\delta^2 - 2r_x\beta^2\sigma_e^2 + r_x^2\beta^2(\sigma_e^2 + \sigma_{x^*}^2) \\
&= \sigma_\delta^2 - 2r_x\beta^2\sigma_e^2 + r_x\beta^2\sigma_e^2 \\
&= \sigma_\delta^2 - r_x\beta^2\sigma_e^2. \qquad (A.21)
\end{aligned}$$

Using the result in Equation (A.21) for the R-squared of Equation (A.19) we have

$$plim\ R_x^2 = 1 - \frac{plim\ \widehat{\sigma}_\delta^2}{plim\ \widehat{\sigma}_{y^*}^2} = 1 - \frac{\sigma_\delta^2 - r_x \beta^2 \sigma_e^2}{\sigma_{y^*}^2}. \tag{A.22}$$

Asymptotically, the ratio of the R-squareds of the model without any measurement error in Equation (A.1) and the model with measurement error in the explanatory variable in Equation (A.19) is given by (using also that $\sigma_{y^*}^2 = \beta^2 \sigma_{x^*}^2 + \sigma_u^2$)

$$\begin{aligned} plim\ \frac{R^{*2}}{R_x^2} &= \frac{(\sigma_{y^*}^2 - \sigma_u^2)/\sigma_{y^*}^2}{(\sigma_{y^*}^2 - \sigma_\delta^2 + r_x \beta^2 \sigma_e^2)/\sigma_{y^*}^2} \\ &= \frac{\sigma_{y^*}^2 - \sigma_u^2}{\sigma_{y^*}^2 - \sigma_u^2 - \beta^2 \sigma_e^2 + r_x \beta^2 \sigma_e^2} \\ &= \frac{\beta^2 \sigma_{x^*}^2}{\beta^2 \sigma_{x^*}^2 - \beta^2 \sigma_e^2 + r_x \beta^2 \sigma_e^2} \\ &= \frac{1}{1 - \frac{\sigma_e^2}{\sigma_{x^*}^2}(1 - r_x)} \\ &= \frac{1}{1 - r_x} \\ &= 1 + \frac{\sigma_e^2}{\sigma_{x^*}^2}. \end{aligned} \tag{A.23}$$

Thus,

$$plim\ R^{*2} = \left(1 + \frac{\sigma_e^2}{\sigma_{x^*}^2}\right) \cdot plim\ R_x^2. \tag{A.24}$$

A.2.3 Measurement Error in both y_i^* and x_i^*

Suppose we want to estimate Equation (A.1) but we have inaccurate measures of both the dependent and the explanatory variable, that is

$$y_i = y_i^* + w_i, \quad Var(w_i) = \sigma_w^2, \ \ E(w_i) = E(w_i y_i^*) = E(w_i u_i) = 0, \tag{A.25}$$

$$x_i = x_i^* + e_i, \quad Var(e_i) = \sigma_e^2, \ \ E(e_i) = E(e_i x_i^*) = E(e_i u_i) = 0, \tag{A.26}$$

Instead of Equation (A.1), the estimated equation is given by

$$\begin{aligned} y_i &= \alpha + \beta \cdot x_i + u_i - \beta \cdot e_i + w_i \\ &= \alpha + \beta \cdot x_i + \delta_i, \end{aligned} \tag{A.27}$$

with $Var(\delta_i) \equiv \sigma_\delta^2 = \beta^2\sigma_e^2 + \sigma_u^2 + \sigma_w^2$. The R-squared for the model with measurement error in the explanatory variable in Equation (A.27) is given by

$$R_{xy}^2 \equiv 1 - \frac{\widehat{\sigma}_\delta^2}{\widehat{\sigma}_y^2}. \tag{A.28}$$

Suppose the measurement errors of the dependent and the explanatory variables are uncorrelated, $E(e_i w_i) = 0$.[28] Then, as measurement error in the dependent variable does not introduce any inconsistency but only increases the variance of the regression error, σ_δ^2, the effects of measurement error in the dependent and explanatory variable on the goodness of fit are independent of each other. Consequently, asymptotically it holds that the ratio of the R-squareds of the model without any measurement error in Equation (A.1) and the model with measurement error in both the dependent and the explanatory variable in Equation (A.27) is given by

$$plim\ \frac{R^{*2}}{R_{xy}^2} = \left(1 + \frac{\sigma_e^2}{\sigma_{x^*}^2}\right) \cdot \left(1 + \frac{\sigma_w^2}{\sigma_{y^*}^2}\right), \tag{A.29}$$

so that

$$plim\ R^{*2} = \left(1 + \frac{\sigma_e^2}{\sigma_{x^*}^2}\right) \cdot \left(1 + \frac{\sigma_w^2}{\sigma_{y^*}^2}\right) \cdot plim\ R_{xy}^2. \tag{A.30}$$

Notice that the this does not imply that for sufficiently small signal-to-noise ratios the true R-squared is larger than unity, as lower signal-to-noise ratios imply lower values of R_{xy}^2. To see this, for the model with measurement error in the dependent variable notice that $R_y^2 \to 0$ for the case when $\frac{\sigma_w^2}{\sigma_{x^*}^2} \to \infty$ since Equation (A.15) can be rewritten as

$$plim\ R_y^2 = 1 - \frac{1 + \frac{\sigma_u^2}{\sigma_w^2}}{1 + \frac{\sigma_u^2}{\sigma_w^2} + \beta^2 \frac{\sigma_{x^*}^2}{\sigma_w^2}}. \tag{A.31}$$

For the model with measurement error in the explanatory variable, observe that upon substitution, Equation (A.22) can be rewritten as

$$plim\ R_x^2 = 1 - \frac{\frac{\sigma_u^2}{\sigma_e^2} + (1 - r_x)\beta^2}{\frac{\sigma_u^2}{\sigma_e^2} + \beta^2 \frac{\sigma_{x^*}^2}{\sigma_e^2}}. \tag{A.32}$$

As measurement error becomes stronger, that is as $\frac{\sigma_{x^*}^2}{\sigma_e^2} \to 0$ implying $r_x \to 1$, for fixed β it holds that $R_x^2 \to 0$.

[28] If they are correlated, this only affects the magnitude of the error variance, σ_δ^2, but leaves the derivations of the effects of measurement error on the R-squared unchanged.

B Constructing Generalized Impulse Response Functions with Monte Carlo Integration

Following Koop et al. (1996), I integrate out the history dependence of the impulse responses implied by the PCHVAR model while preserving the cross-sectional spread in the impulse responses by repeating the following steps in country-specific Monte Carlo experiments:

1. Pick an initial observation $\tilde{\boldsymbol{z}}_{i1}$ by drawing randomly from the actual data $\{\boldsymbol{z}_{it}\}_{t=1,2,\ldots,T_i}$ of country i.

2. Generate a simulated time series $\{\tilde{\boldsymbol{z}}_{it}\}_{t=1,2,\ldots,H}$ of the conditioning variable of country i by iterating on
$$\tilde{z}_{ijt} = \widehat{\phi}_0 + \widehat{\phi}_1 \cdot \tilde{z}_{ij,t-1} + u_{ijt}, \tag{B.1}$$
for $j = 1, 2, 3$ and where $u_{ijt} \overset{i.i.d.}{\sim} N(0, \widehat{\sigma}^2_{ij})$ and $\widehat{\phi}_0$, $\widehat{\phi}_1, \widehat{\sigma}^2_{ij}$ are OLS estimates.[29]

3. Calculate history-dependent impulse responses using the simulated values of the conditioning variable obtained in steps 1. and 2.

4. Repeat steps 1. to 3. 250 times and store the impulse responses in each replication.

5. Calculate the average of the impulse responses obtained in step 4.

The intuition for this approach is the following: The impulse response at horizon h

$$\boldsymbol{ir}^{(PCHVAR)}(h) \quad \equiv \quad E\big[\boldsymbol{y}_{i,t+h}|\boldsymbol{u}_{it}, \boldsymbol{Y}_{it}, \underline{\boldsymbol{\mathcal{Z}}}^{(h)}_{i,t+h}\big] - E\big[\boldsymbol{y}_{i,t+h}|\boldsymbol{Y}_{it}, \underline{\boldsymbol{\mathcal{Z}}}^{(h)}_{i,t+h}\big], \tag{B.2}$$

is a random variable because $\underline{\boldsymbol{\mathcal{Z}}}^{(h)}_{i,t+h}$ is a random variable (for example countries' structural characteristics) of which $\boldsymbol{\mathcal{Z}}^{(h)}_{i,t+h}$ is a realization, and $\boldsymbol{Y}_{it} \equiv [\boldsymbol{y}_{it}, \boldsymbol{y}_{i,t-1}, \ldots, \boldsymbol{y}_{i,t-p}]$. To integrate out the history dependence, calculate

$$E\left[ir^{(PCHVAR)}(h, \boldsymbol{u}_{it}, \boldsymbol{Y}_{it}, \underline{\boldsymbol{\mathcal{Z}}}^{(h)}_{i,t+h})\right], \tag{B.3}$$

where the expectation is taken with respect to $\underline{\boldsymbol{\mathcal{Z}}}^{(h)}_{i,t+h}$.

[29] I smooth the simulated time series $\{\tilde{\boldsymbol{z}}_{it}\}_{t=1,2,\ldots,H}$ using the Hodrick-Prescott filter before moving to step 3. Moreover, I drop a replication if the simulated time series exceeds unity or declines below zero, as in the estimation of the PCHVAR model re-scaled structural characteristics that fall in $[0, 1]$ are used, see Section 4.4.

C Tables

Table 1: Cross-Country Rankings of Statistics of the Responses of Output to a Monetary Policy Shock from Individual Country VAR Models

Maximum Response		Mean Response		Response at Horizon $H = 48$	
Poland	-0.003	Poland	-0.002	Italy	0.001
Czech Republic	-0.004	Czech Republic	-0.002	Germany	-0.000
Hungary	-0.005	United States	-0.002	United States	-0.000
Australia	-0.005	Australia	-0.003	Hungary	-0.001
United Kingdom	-0.005	Hungary	-0.003	Czech Republic	-0.002
Canada	-0.005	United Kingdom	-0.004	Poland	-0.002
New Zealand	-0.005	New Zealand	-0.004	Australia	-0.003
United States	-0.006	Canada	-0.004	Denmark	-0.003
Germany	-0.008	Germany	-0.004	France	-0.004
Switzerland	-0.008	Switzerland	-0.005	New Zealand	-0.004
France	-0.009	Italy	-0.005	United Kingdom	-0.005
Italy	-0.011	France	-0.006	Canada	-0.005
Portugal	-0.012	Portugal	-0.008	Portugal	-0.005
Denmark	-0.014	Denmark	-0.009	Switzerland	-0.007
Sweden	-0.016	Korea	-0.009	Korea	-0.007
Austria	-0.016	Sweden	-0.011	Belgium	-0.010
Spain	-0.016	Spain	-0.011	Spain	-0.012
Korea	-0.017	Belgium	-0.011	Sweden	-0.012
Belgium	-0.017	Austria	-0.011	Austria	-0.012
Ireland	-0.019	Ireland	-0.013	Ireland	-0.014
Mean	-0.010	Mean	-0.006	Mean	-0.005
Std.	0.005	Std.	0.004	Std.	0.004

Note: The table provides the country rankings of the maximum, mean and the values at horizon $H = 48$ of the responses of output to a monetary policy shock derived from the country VAR models in Equation (1). For each of these statistic, the table also provides the mean and the standard deviation across countries.

Table 2: Cross-Country Rankings of Statistics of the Responses of Prices to a Monetary Policy Shock from Individual Country VAR Models

Maximum Response		Mean Response		Response at Horizon $H = 48$	
New Zealand	-0.002	Australia	0.000	New Zealand	-0.002
Australia	-0.002	Sweden	-0.000	Australia	-0.002
Sweden	-0.002	New Zealand	-0.001	Sweden	-0.002
Switzerland	-0.003	Italy	-0.001	Poland	-0.003
Poland	-0.003	Poland	-0.001	United States	-0.003
United Kingdom	-0.003	Switzerland	-0.001	Switzerland	-0.003
United States	-0.003	France	-0.001	United Kingdom	-0.003
Hungary	-0.003	United States	-0.001	Czech Republic	-0.003
Germany	-0.004	United Kingdom	-0.002	Hungary	-0.003
France	-0.004	Hungary	-0.002	Canada	-0.003
Czech Republic	-0.004	Germany	-0.002	France	-0.004
Korea	-0.005	Austria	-0.003	Germany	-0.004
Austria	-0.005	Ireland	-0.003	Korea	-0.004
Italy	-0.005	Korea	-0.003	Austria	-0.005
Denmark	-0.007	Czech Republic	-0.003	Italy	-0.005
Ireland	-0.007	Canada	-0.004	Denmark	-0.006
Canada	-0.008	Denmark	-0.005	Ireland	-0.007
Portugal	-0.008	Portugal	-0.005	Portugal	-0.008
Spain	-0.010	Spain	-0.006	Belgium	-0.010
Belgium	-0.010	Belgium	-0.007	Spain	-0.010
Mean	-0.005	Mean	-0.002	Mean	-0.004
Std.	0.003	Std.	0.002	Std.	0.003

Note: The table provides the country rankings of the maximum, mean and the values at horizon $H = 48$ of the responses of prices to a monetary policy shock derived from the country VAR models in Equation (1). For each of these statistic, the table also provides the mean and the standard deviation across countries.

Table 3: Cross-Country Differences in Structural Characteristics

Country	Financial Structure Index			Strictness of Employment Protection			Share of Manufacturing Durable Components		
	Mean	Max	Min	Mean	Max	Min	Mean	Max	Min
Australia	0.08	0.44	-0.33	1.16	1.20	0.91	0.06	0.07	0.04
Austria	0.13	0.21	-0.03	2.09	2.22	1.92	0.12	0.13	0.11
Belgium	0.11	0.21	-0.02	2.33	3.33	2.13	0.08	0.09	0.06
Canada	0.21	0.63	-0.02	0.75	0.75	0.75	0.08	0.10	0.06
Czech Republic	-0.26	-0.10	-0.33	1.94	2.04	1.89	0.16	0.17	0.15
Denmark	0.15	1.65	-0.80	1.53	1.98	1.48	0.08	0.08	0.07
France	0.13	0.38	-0.03	3.01	3.06	2.96	0.08	0.09	0.06
Germany	0.18	0.38	-0.13	2.20	2.35	2.10	0.15	0.15	0.14
Hungary	-0.83	-0.28	-1.50	1.40	1.67	1.26	0.11	0.13	0.08
Ireland	0.72	1.44	-0.00	1.01	1.12	0.92	0.09	0.11	0.06
Italy	-0.11	0.43	-0.28	2.41	3.67	1.80	0.11	0.11	0.09
Korea	0.46	0.69	-0.01	2.18	2.80	1.83	0.17	0.20	0.16
New Zealand	0.25	0.81	-1.52	1.24	1.49	0.84	0.05	0.06	0.05
Poland	-0.87	-0.39	-1.22	1.58	1.92	1.35	0.09	0.10	0.08
Portugal	0.43	0.92	-0.12	3.58	3.87	3.00	0.07	0.08	0.06
Spain	0.16	0.90	-0.08	2.98	3.02	2.93	0.09	0.09	0.07
Sweden	0.36	0.90	-0.05	2.23	2.51	1.67	0.12	0.13	0.10
Switzerland	0.40	0.53	0.04	1.14	1.14	1.14	0.11	0.12	0.11
United Kingdom	0.39	0.69	0.08	0.69	0.75	0.60	0.08	0.10	0.05
United States	0.15	0.35	-0.09	0.21	0.21	0.21	0.06	0.08	0.06
Mean	0.11	0.54	-0.32	1.78	2.05	1.59	0.10	0.11	0.08
Std.	0.39	0.50	0.51	0.86	1.03	0.79	0.03	0.03	0.03
Max	0.72	1.65	0.08	3.58	3.87	3.00	0.17	0.20	0.16
Min	-0.87	-0.39	-1.52	0.21	0.21	0.21	0.05	0.06	0.04

Note: The table provides the mean, maximum and minimum values of the financial structure index, Strictness of Employment Protection Indicator, and the share of durable goods manufacturing for every country in the sample and the time periods displayed in Table 11. The table also provides the mean, maximum and minimum values as well as the standard deviation across all countries of the country-specific figures.

Table 4: Regressing Statistics of Country VAR Model Impulse Responses of Output on Countries' Structural Characteristics

	$f(\cdot)$		
	max. IR	mean IR	IR at H
Financial Structure Index	-0.007^{***} (0.002)	-0.005^{***} (0.001)	-0.006^{***} (0.002)
Strictness of Employment Protection	-0.002^{**} (0.001)	-0.002^{**} (0.001)	-0.001 (0.001)
Share of Manufacturing Durable Components	-0.019 (0.026)	-0.004 (0.017)	-0.006 (0.022)
R^2	0.47	0.47	0.30

Note: The table displays the estimates $\widehat{b}$ from Equation (2). Standard errors of point estimates are reported in parentheses. * (**, ***) represents statistical significance at the 10% (5%, 1%) significance level. Standard errors are heteroskedasticity-consistent.

Table 5: Regressing Statistics of Country VAR Model Impulse Responses of Prices on Countries' Structural Characteristics

	$f(\cdot)$		
	max. IR	mean IR	IR at H
Financial Structure Index	-0.001^{*} (0.001)	-0.001 (0.000)	-0.001^{*} (0.001)
Strictness of Employment Protection	-0.001^{*} (0.001)	-0.001^{*} (0.000)	-0.002^{***} (0.001)
Share of Manufacturing Durable Components	0.013 (0.013)	0.004 (0.010)	0.018 (0.011)
R^2	0.24	0.15	0.40

Note: The table displays the estimates $\widehat{b}$ from Equation (2). Standard errors of point estimates are reported in parentheses. * (**, ***) represents statistical significance at the 10% (5%, 1%) significance level. Standard errors are heteroskedasticity-consistent.

Table 6: Wald Test for $H_0 : \boldsymbol{ir}_i^{(VAR)} = \boldsymbol{ir}_i^{(PCHVAR)}$

Country	Output	Prices
Australia		
Austria		
Belgium		
Canada	***	***
Czech Republic	**	
Denmark		**
France		
Germany		
Hungary		
Ireland		
Italy		**
Korea		
New Zealand		
Poland		
Portugal		**
Spain		***
Sweden		
Switzerland		***
United Kingdom		
United States		

Note: The table reports the results for Wald tests with the null that the impulse responses from the country VAR models are identical with those from the PCHVAR model. ***, **, * represent rejection of the null at the 1, 5 and 10 percent significance level. The variance matrices and the critical values are obtained from a bootstrap.

Table 7: Rank Correlations Between Statistics of the Country VAR Model and the PCHVAR Model Impulse Responses

	Output	Prices
Maximum Response	0.91***	0.59***
Mean Response	0.88***	0.31
Response at H	0.70***	0.72***

Note: The table reports rank correlations between impulse response statistics from the country VAR models and the PCHVAR model. * (**, ***) represents statistical significance at the 10% (5%, 1%) significance level.

Table 8: Correlation of Country VAR Model Impulse Responses Between Baseline and Alternative Specifications

Robustness Check	Max. IR	IR at H=48	Mean IR	All Horizons
Exchange Rate	0.90, 0.92	0.88, 0.89	0.89, 0.95	0.95, 0.95
Money	0.87, 0.71	0.31, 0.41	0.70, 0.86	0.83, 0.84
Overnight Rate	0.93, 0.93	0.97, 0.92	0.96, 0.96	0.97, 0.96
p=9	0.93, 0.88	0.92, 0.88	0.91, 0.88	0.94, 0.90
q=3	0.90, 0.87	0.93, 0.87	0.87, 0.84	0.96, 0.90
Euro Area Aggregates	0.77, 0.59	0.50, 0.37	0.71, 0.51	0.80, 0.81
Financial Structure	0.93, 0.76	0.75, 0.69	0.90, 0.82	0.85, 0.81
Share of Manufacturing Durable Components	0.73, 0.79	0.76, 0.67	0.74, 0.84	0.93, 0.84
Strictness of Employment Protection	0.87, 0.86	0.86, 0.78	0.81, 0.94	0.89, 0.86

Note: The table reports the correlations between the impulse responses of the baseline country VAR model and the robustness checks described in Section 6. For each robustness check and each impulse response statistic, the first correlation refers to the impulse responses of output and the second to those of prices.

Table 9: Monetary Policy Operating Frameworks

Country	Inflation Target	Policy Rate/Operating Target	Operating Framework	Comments
Australia: Reserve Bank of Australia (RBA).	Full-fledged. Target is 2 to 3% in the medium term.	Target for Cash Rate/Cash Rate (overnight money market rate).	The RBA uses open market operations on a day-to-day basis to keep the Cash Rate close to target. The RBA offers a standing marginal lending facility 25 basis points above the Cash Rate.	-
Canada: Bank of Canada (BoC).	Full-fledged. Target is 1 to 3% over the next six to eight quarters.	Bank Rate/overnight money market rate.	The BoC maintains an interest rate corridor. The upper bound is the Bank Rate, the rate for the standing marginal lending facility; the implicit target for the money market rate is 25 basis points below the Bank rate, and 25 basis points above the rate of the standing deposit facility. The BoC uses open market operations on a day-to-day basis to keep the money market rate close to target.	-
Czech Republic: Czech National Bank (CNB).	Full-fledged. The target is 2% (since 2005).	Repo Rate/overnight money market rate.	The CNB maintains an interest rate corridor. The CNB conducts repo auctions every two weeks.	Inflation targeting was introduced in 1998. Before 2001, the CNB targeted net inflation rather than consumer price inflation. The inflation targets have varied over time, but have been decreasing monotonically.
Denmark: Danmark Nationalbank (DNB).	Eclectic. Exchange rate target.	Lending Rate/overnight money market rate.	The DNB conducts weekly repo auctions in which it grants seven-day loans to counterparties at the Lending Rate. The overnight money market rate is to be close to the Lending Rate. Counterparties can also deposit funds overnight at the Current-Account Rate, but may not lend funds overnight. In case of liquidity shortage, the DNB intervenes by open market operations.	Denmark maintains a fixed-exchange-rate policy vis-à-vis the Euro Area.
Euro Area: European Central Bank (ECB).	Eclectic. Price stability is defined as a year-on-year increase in Euro Area consumer prices of below 2%.	Repo Rate/overnight money market rate.	The ECB maintains an interest rate corridor. The ECB conducts weekly repo auctions.	-
Hungary: Magyar National Bank (MNB)	Full-fledged. The target is 3% (since 2007) in the medium term.	MNB-Bill Rate/three-month money market rate.	The MNB maintains an interest rate corridor. The MNB conducts weekly repo auctions.	From 1995 to 2007, the Forint was tied to a crawling basket of currencies (US$, DM and Euro, respectively; only Euro since 2000). Since 2008, the Forint is floating vis-à-vis the Euro. Inflation targeting was adopted in 2001. The inflation target has been varying.
South Korea: Bank of Korea (BoK).	Full-fledged. The target is 3% over the next three years.	BoK Base Rate/overnight money market rate.	The BoK maintains an interest rate corridor. The BoK conducts weekly repo auctions.	Inflation targeting was adopted in 1998.

Continued on next page

Table 9: Monetary Policy Operating Frameworks (continued)

Country	Inflation Target	Policy Rate/Operating Target	Operating Framework	Comments
New Zealand: Reserve Bank of New Zealand (RBNZ).	Full-fledged. The target is 1 to 3% (since 2008) over the medium term.	Official Cash Rate/overnight money market rate.	The RBN maintains an interest rate corridor. The RBN conducts open market operations on a day-by-day basis to keep the overnight money market rate close to the Official Cash Rate.	The RBN switched to targeting the overnight money market rate instead of settlement cash balances in 1999.
Poland: National Bank of Poland (NBP).	Full-fledged. The target is 2.5% (since 2004) for the next year.	Reference Rate/overnight money market rate.	The NBP maintains an interest rate corridor. The NBP conducts weekly repo auctions.	Inflation targeting was adopted in 1999.
Sweden: Sveriges Riksbank (SRB).	Full-fledged. The target is 2%.	Repo Rate/overnight money market rate.	The SRB maintains an interest rate corridor. The SRB conducts weekly repo auctions.	-
Switzerland: Swiss National Bank (SNB).	Eclectic. The SNB equates price stability with a rise in consumer prices of less than 2% per year.	Repo Rate/three-month money market rate.	The SNB maintains an interest rate corridor. The SNB conducts weekly repo auctions.	-
United Kingdom: Bank of England (BoE).	Full-fledged. The target is 2%.	BoE Base Rate/overnight money market rate.	The BoE maintains an interest rate corridor. The BoE conducts weekly repo auctions.	The operational framework of the BoE was reformed in 2006.
United States: Federal Reserve System (Fed).	Eclectic. The Fed is to promote maximum sustainable output and employment and to promote stable prices, with no explicit inflation targeting.	Fed Funds Rate target/overnight money market rate.	The Fed conducts open market operations on a day-by-day basis to keep the Fed Funds Rate close to its target. The Fed also provides a standing marginal lending facility.	-

Note: Central banks with an eclectic inflation-targeting regime feature an implicit price stability anchor and are not fully transparent and accountable with respect to an inflation target. In contrast, central banks with full-fledged inflation targeting clearly commit to their inflation target, and institutionalize this commitment in the form of a transparent monetary framework that fosters accountability of the central bank to the target. See Carare and Stone (2006).

Table 10: An Overview of the Literature on Cross-Country Asymmetries in Monetary Transmission

Study	Approach	Countries	Finding
Gerlach and Smets (1995)	country-specific VAR models with mix of long- and short-run restrictions; one standard deviation shock and 100 basis points shock over eight quarters	CAN, FRA, DEU, ITA, JPN, GBR, USA	"The estimates of the effects of monetary policy provide little evidence of large differences in the monetary transmission between countries, particularly not when estimated confidence bands are taken into account." (p. 39)
Barran et al. (1996)	identical, country-specific VAR models with recursive identification; one standard deviation shock	AUT, DEU, DNK, ESP, FIN, FRA, ITA, NLD, GBR	"[European countries] were similar in the sense of responses and lags [to a monetary policy shock]. However, the magnitudes involved seem different." (p. 21)
Ramaswamy and Sløk (1998)	country-specific VAR models with recursive identification; 100 basis points shock	AUT, BEL, DEU, DNK, ESP, FIN, FRA, GBR, ITA, NLD, PRT, SWE	"there appear to be marked differences in the real effects of monetary policy among EU countries." (p. 383)
Dornbusch et al. (1998)	Estimate country-specific monetary policy reaction function and dynamic output equations, so that they can simulate a currency union-wide shock	DEU, ESP, FRA, ITA, SWE, GBR	"the impact effect on output is always significant, but different across countries" (p. 40)
Kieler and Saarenheimo (1998)	identical, country-specific VAR models; compare impulse responses obtained from rotations of orthogonalized reduced form shocks and impose plausibility windows	DEU, FRA, GBR	"once the uncertainty involved in the structural identification is accounted for, no statistically significant differences in monetary transmission can be found for a group of three large EU countries." (p. 4)
Kim (1999)	identical, country-specific structural VAR models; one standard deviation shock	CAN, DEU, FRA, GBR, ITA, JPN, USA	"the output responses [to a monetary policy shock] are very similar across countries." (p. 399)
Ehrmann (2000)	country-specific and different VAR models; restrictions on the co-integrating properties; ten basis points shock	AUT, BEL, DEU, DNK, FIN, FRA, IRL, ITA, NLD, PRT, ESP, SWE, GBR	"considerable differences in the monetary transmission mechanism" (p. 78)
Clements, Kontolemis and Levy (2001)	country-specific VAR models with recursive identification; 100 basis points shock	AUT, BEL, DEU, FIN, FRA, IRL, ITA, NLD, PRT, ESP	"[the effects of monetary policy] on economic activity are likely to differ across EMU countries" (p. 1)
Mihov (2001)	country-specific VAR models with recursive identification; 100 basis points shock	AUS, AUT, CAN, DEU, FRA, GBR, ITA,JPN, NLD, USA	"monetary policy transmission mechanisms were quite heterogeneous across EMU members." (p. 372)
Mojon and Peersman (2001)	country-specific, structural VAR models; one standard deviation shock	AUT, BEL, DEU, FIN, FRA, GRC, IRL, ITA, NLD, ESP	"the overall pattern for output and prices is quite similar across countries" (p. 17)
Sala (2002)	Dynamic factor model to investigate currency-union wide shock to monetary policy	AUS, BEL, DEU, ESP, FRA, ITA, NLD, PRT	"European countries (...) are characterized by quantitatively different responses [of output to a monetary policy shock]." (p. 1)
Dedola and Lippi (2005)	country-specific VAR models with recursive identification; one standard deviation shock	DEU, FRA, GBR, ITA, USA	"hardly detectable cross-country variability." (p. 1543)
Ciccarelli and Rebucci (2006)	Approach similar to Dornbusch et al. (1998) with reaction functions and output equations, but allow for time variation	DEU, ESP, FRA, ITA	"cross-country differences in the effects of [monetary policy] shocks have not decreased over time." (p. 737)

Note: The table presents an overview of papers on cross-country asymmetries in monetary transmission.

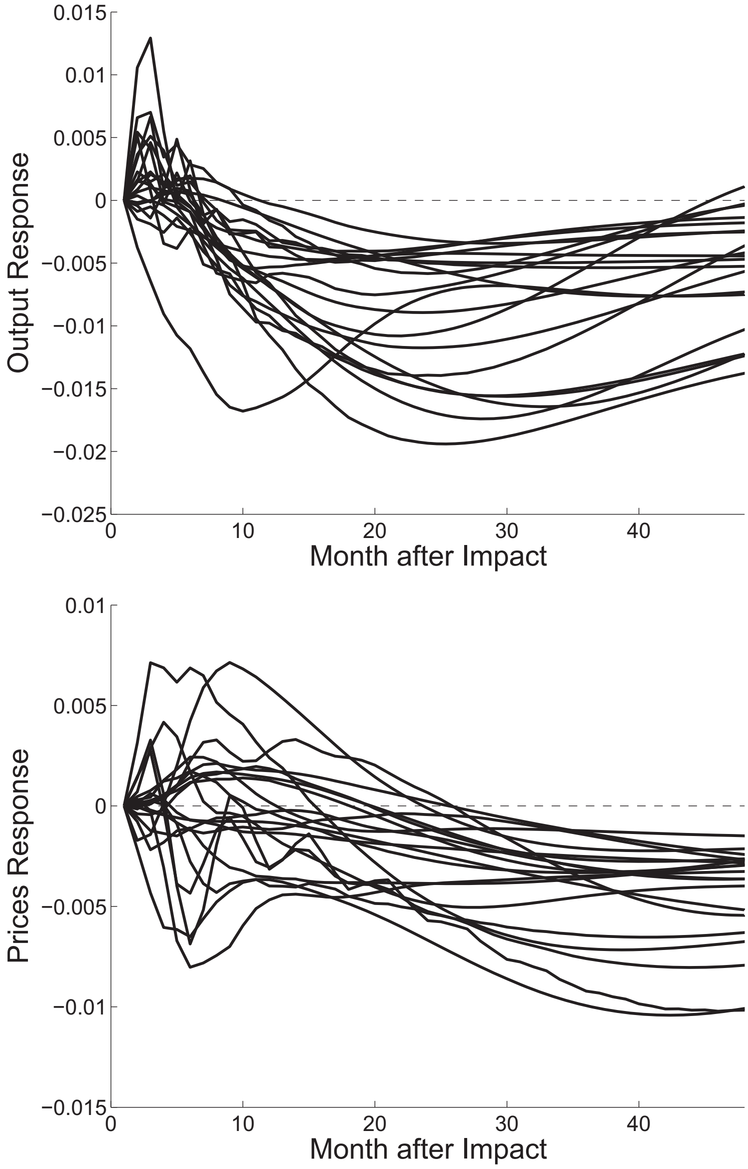
Table 11: Countries Included, Periods Covered and Variables Used

Country	Period Covered	Variables		
Australia	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Austria	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Belgium	1995:1-2009:10	Log CPI (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Canada	1995:1-2009:10	Log CPI (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Czech Republic	1998:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Denmark	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
France	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Germany	1998:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Hungary	1995:3-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Immediate Interest Rate
Ireland	1995:1-2009:10	Log CPI (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Italy	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Korea	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
New Zealand	1995:1-2009:10	Log CPI (nat. sources, q)	Log Real GDP (OECD)	Short-Term Interest Rate
Poland	1995:4-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Portugal	1995:1-2009:10	Log CPI (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Spain	1995:1-2009:10	Log CPI (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
Sweden	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Immediate Interest Rate
Switzerland	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
United Kingdom	1995:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate
United States	1999:1-2009:10	Log GDP Deflator (OECD)	Log Real GDP (OECD)	Short-Term Interest Rate

Note: The table provides a list of the countries included, the time period covered for each country and the variables used. The GDP and deflator figures are taken from the OECD'S Quarterly National Accounts. The CPI figures are taken from the OECD's Prices Database in the Main Economic Indicators and from Statistics New Zealand. The figures are seasonally adjusted. The GDP figures are expressed in US dollars using fixed PPP weights. The short-term interest rate is taken from the Financial Indicators in the OECD's Main Economic Indicators and usually is either the three month interbank offer rate or the rate associated with Treasury bills, Certificates of Deposit or comparable instruments, each of three month maturity. The immediate interest rate is the official discount rate or overnight call-money rate.

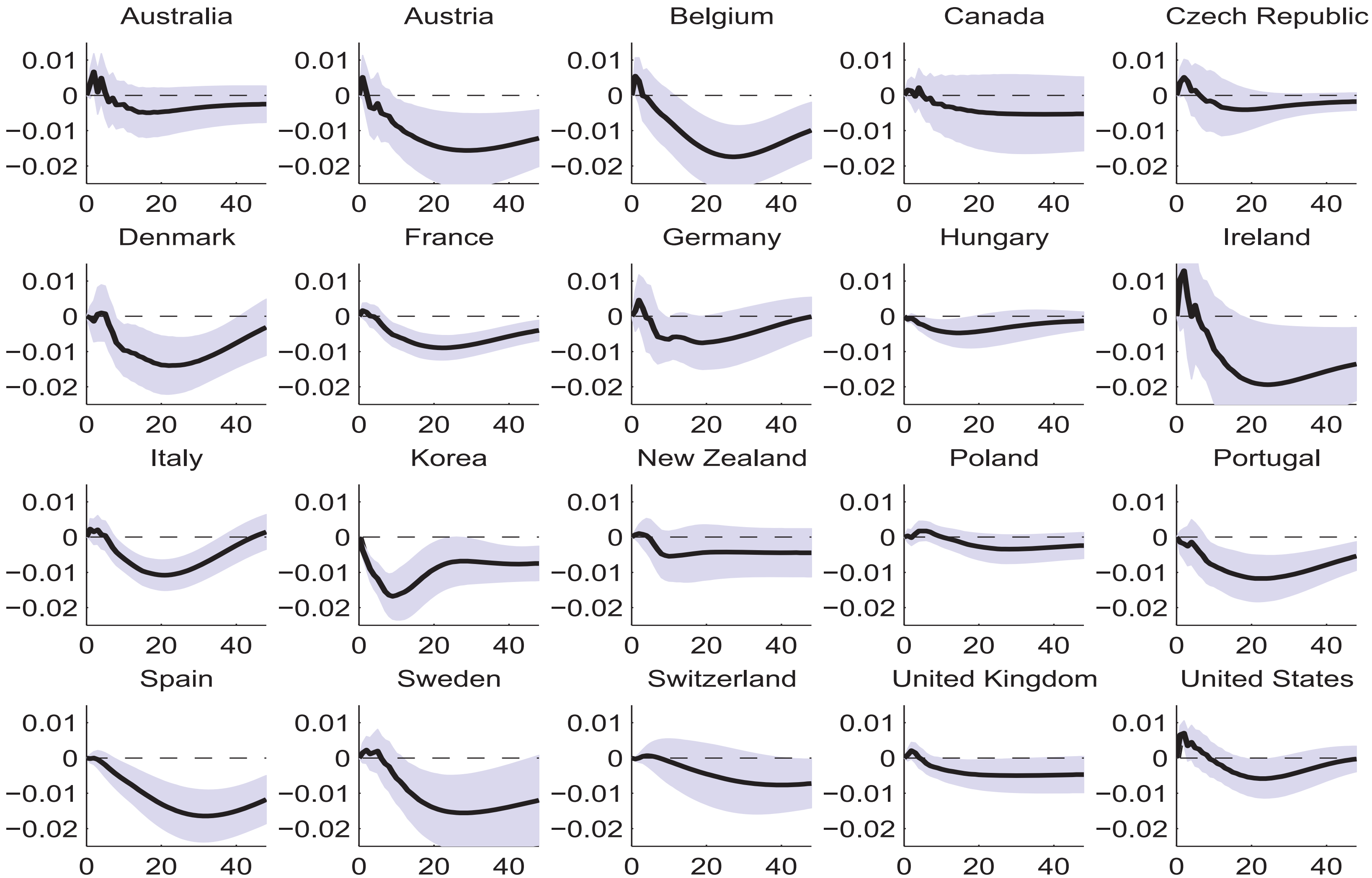
D Figures

Figure 1: Countries' Responses to a Contractionary Monetary Policy Shock



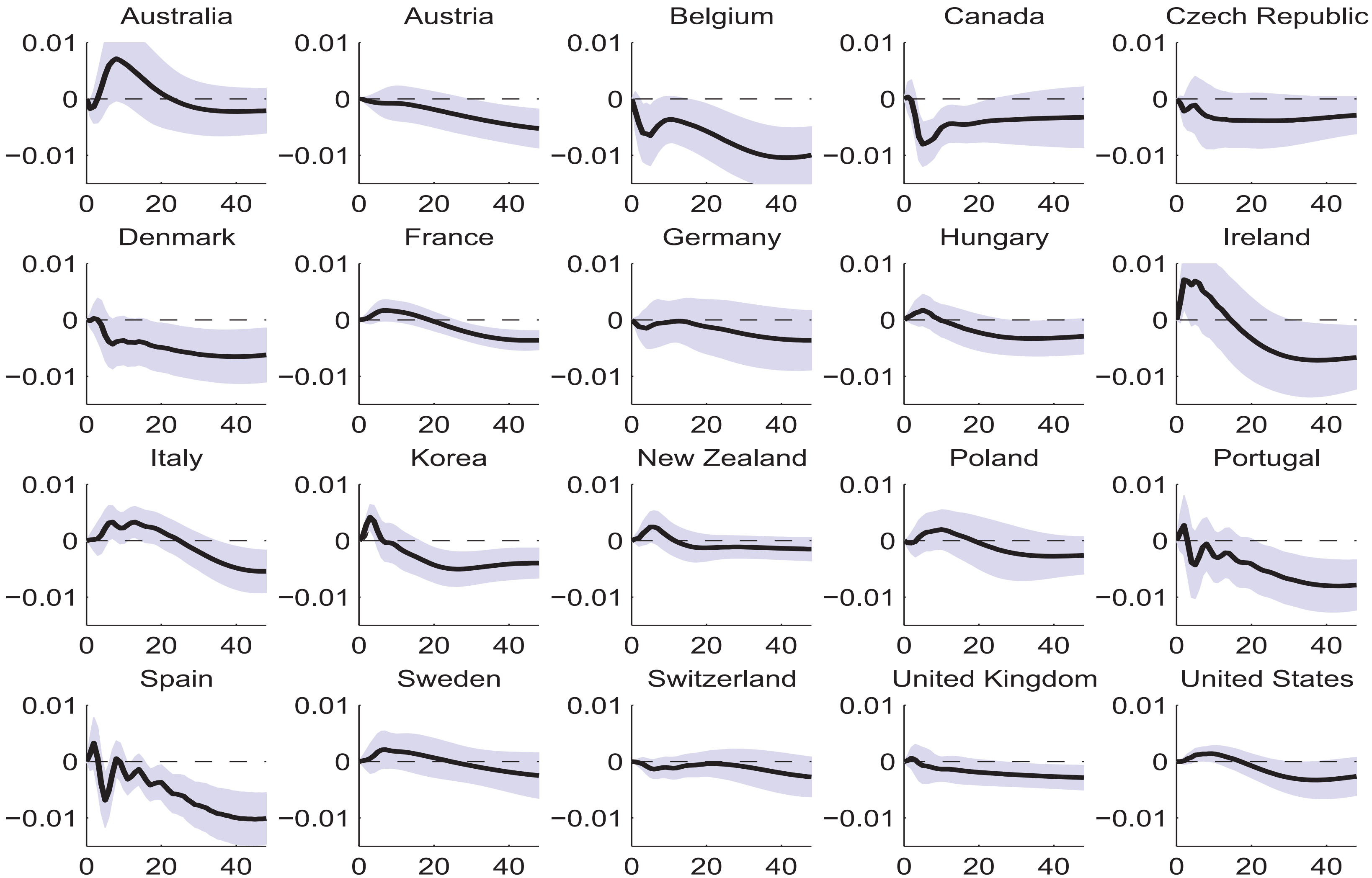
Note: The figure displays the responses of output and prices for all countries in the sample to a 100 basis points monetary policy shock obtained from the country VAR models in Equation (1). The upper panel combines the responses of output for all countries in the sample, and the lower panel combines the responses of prices.

Figure 2: Countries' Output Responses to Contractionary Monetary Policy Shock



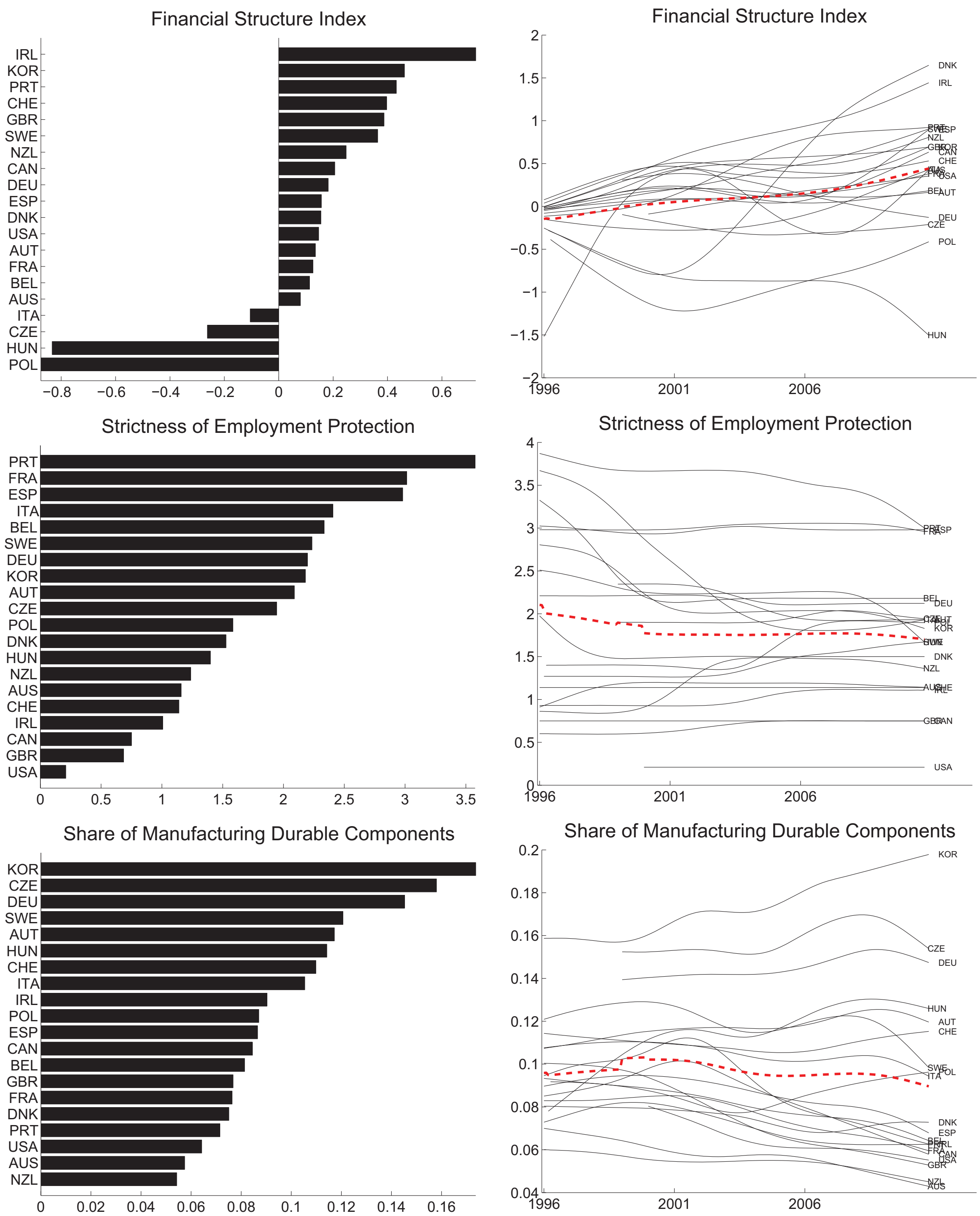
Note: The figure displays the responses of output to a 100 basis points monetary policy shock obtained from the country VAR models in Equation (1). In each panel, the solid line depicts the point estimate of the impulse response and the shaded area represents asymptotic (bootstrap, dark shaded area) 90% confidence bands.

Figure 3: Countries' Price Level Responses to Contractionary Monetary Policy Shock



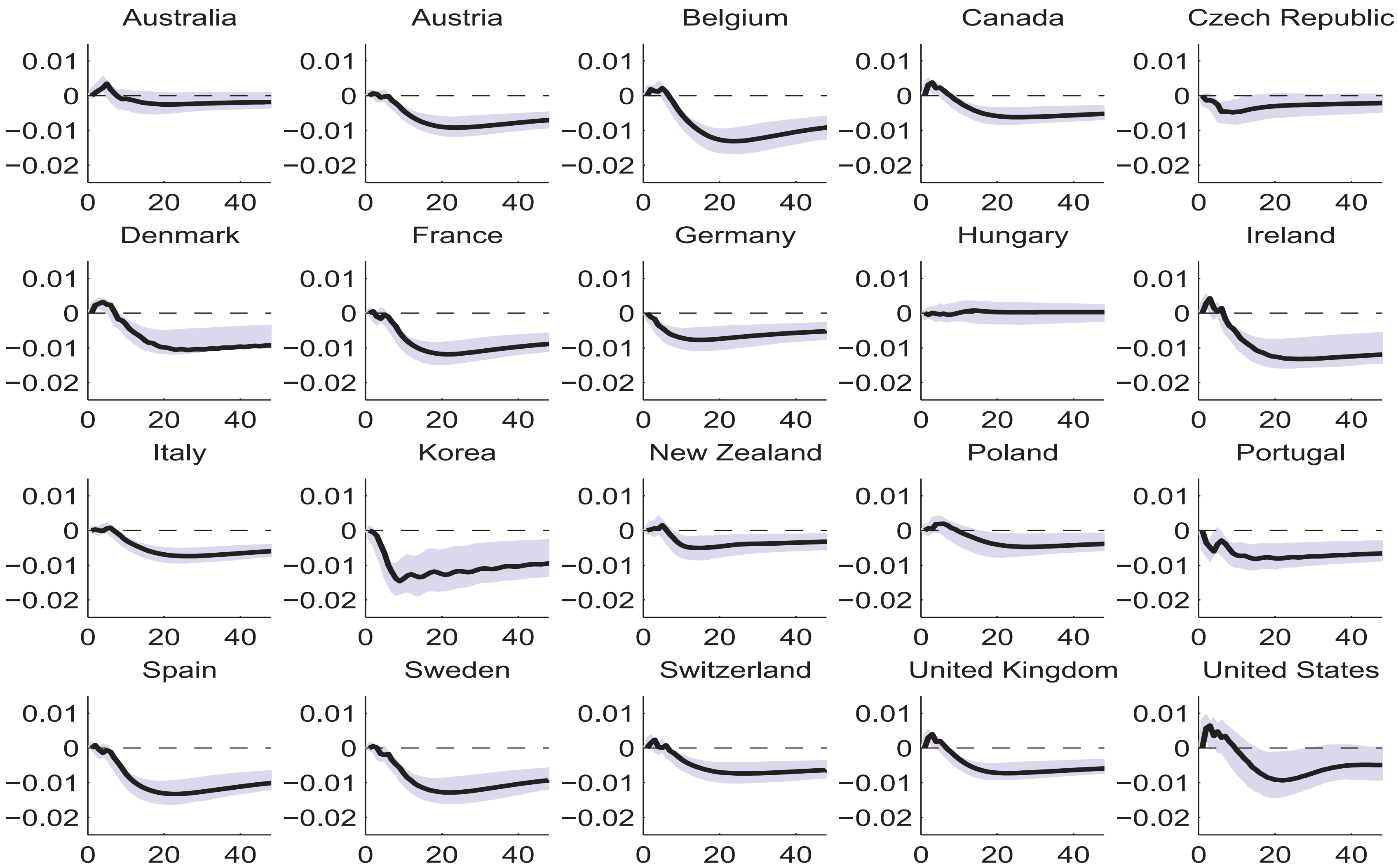
Note: The figure displays the responses of prices to a 100 basis points monetary policy shock obtained from the country VAR models in Equation (1). In each panel, the solid line depicts the point estimate of the impulse response and the shaded area represents asymptotic (bootstrap, dark area) 90% confidence bands.

Figure 4: Cross-Country Differences in Structural Characteristics



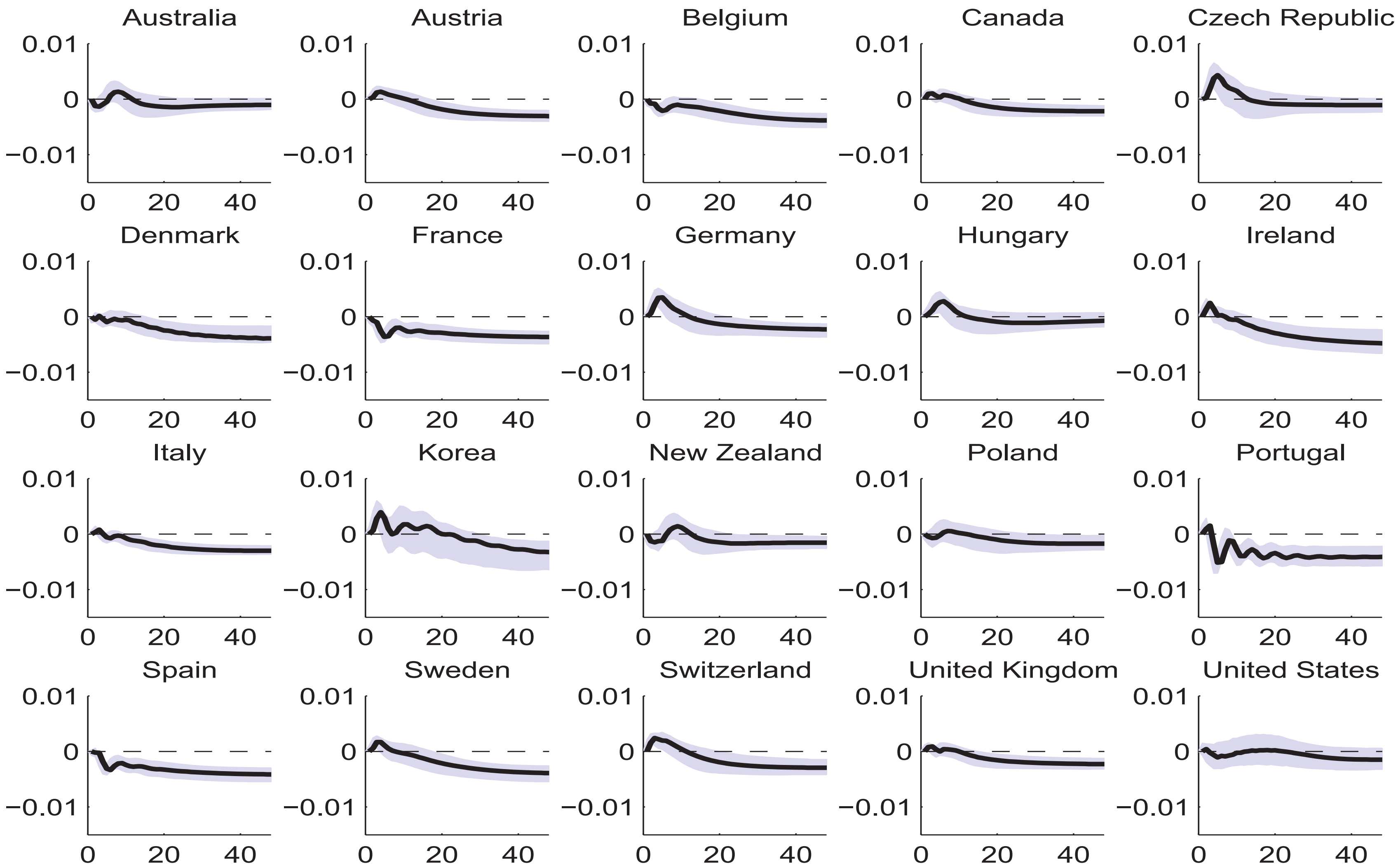
Note: The figure displays the financial structure index, the Strictness of Employment Protection indicator and the share of durable goods manufacturing in total output. In the left-hand side panels, the time averages of the structural characteristics are depicted. In the right-hand side panels, the evolution of the structural characteristics over time is depicted. The solid lines represent the evolution for each country, and the dashed lines the country averages. Larger values of the financial structure index reflect financial systems in which bank credit is more important and in which banking sector competitive pressures are stronger. The Strictness of Employment Protection indicator is bounded between zero and four with larger values reflecting more rigid labor markets.

Figure 5: PCHVAR Model Output Responses to a Contractionary Monetary Policy Shock



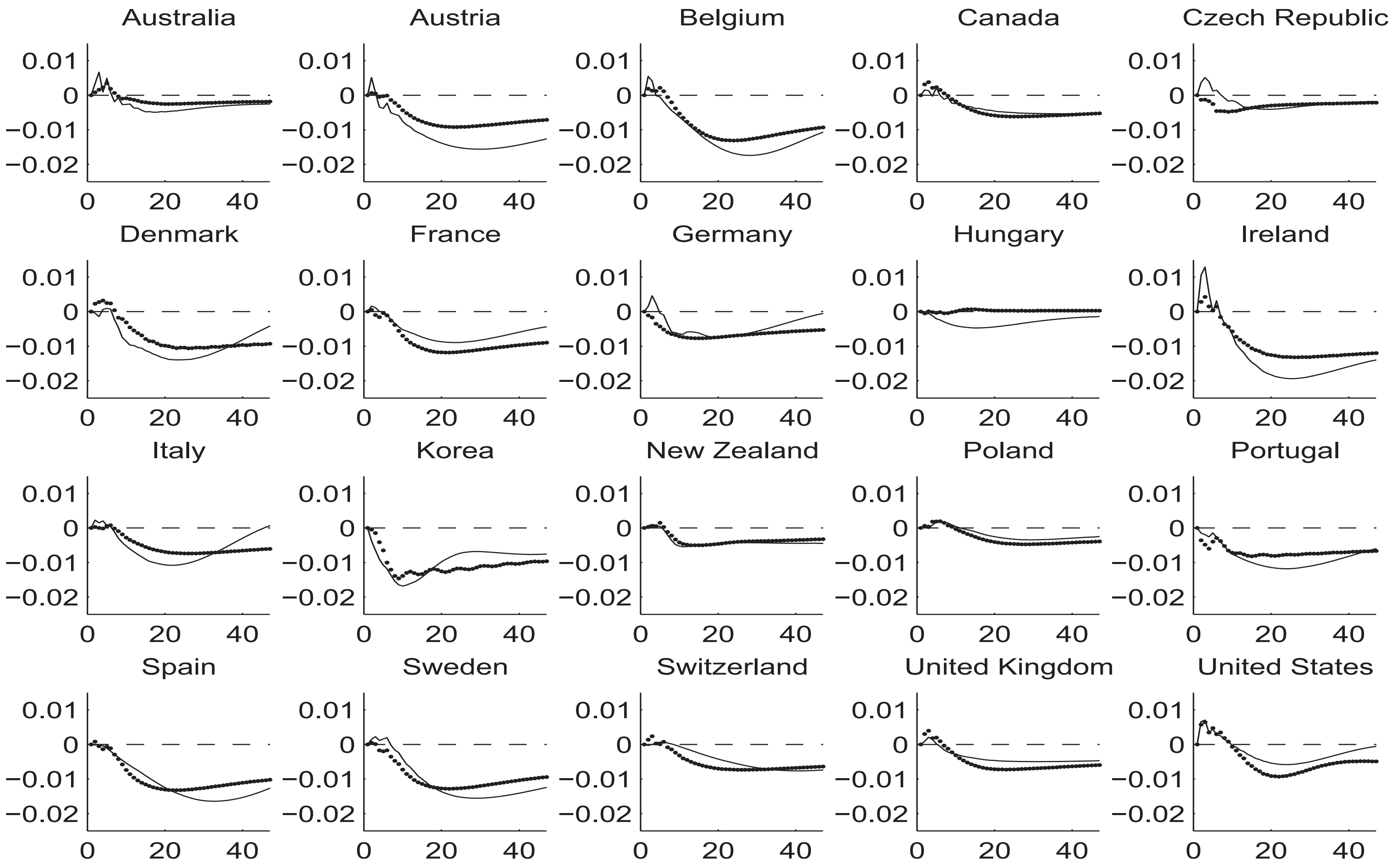
Note: The figure displays the output responses to a 100 basis points monetary policy shock obtained from the the PCHVAR model in Equation (8). In each panel, the solid line depicts the point estimates of the impulse responses and the shaded area 95% bootstrap confidence bands.

Figure 6: PCHVAR Model Price Level Responses to a Contractionary Monetary Policy Shock



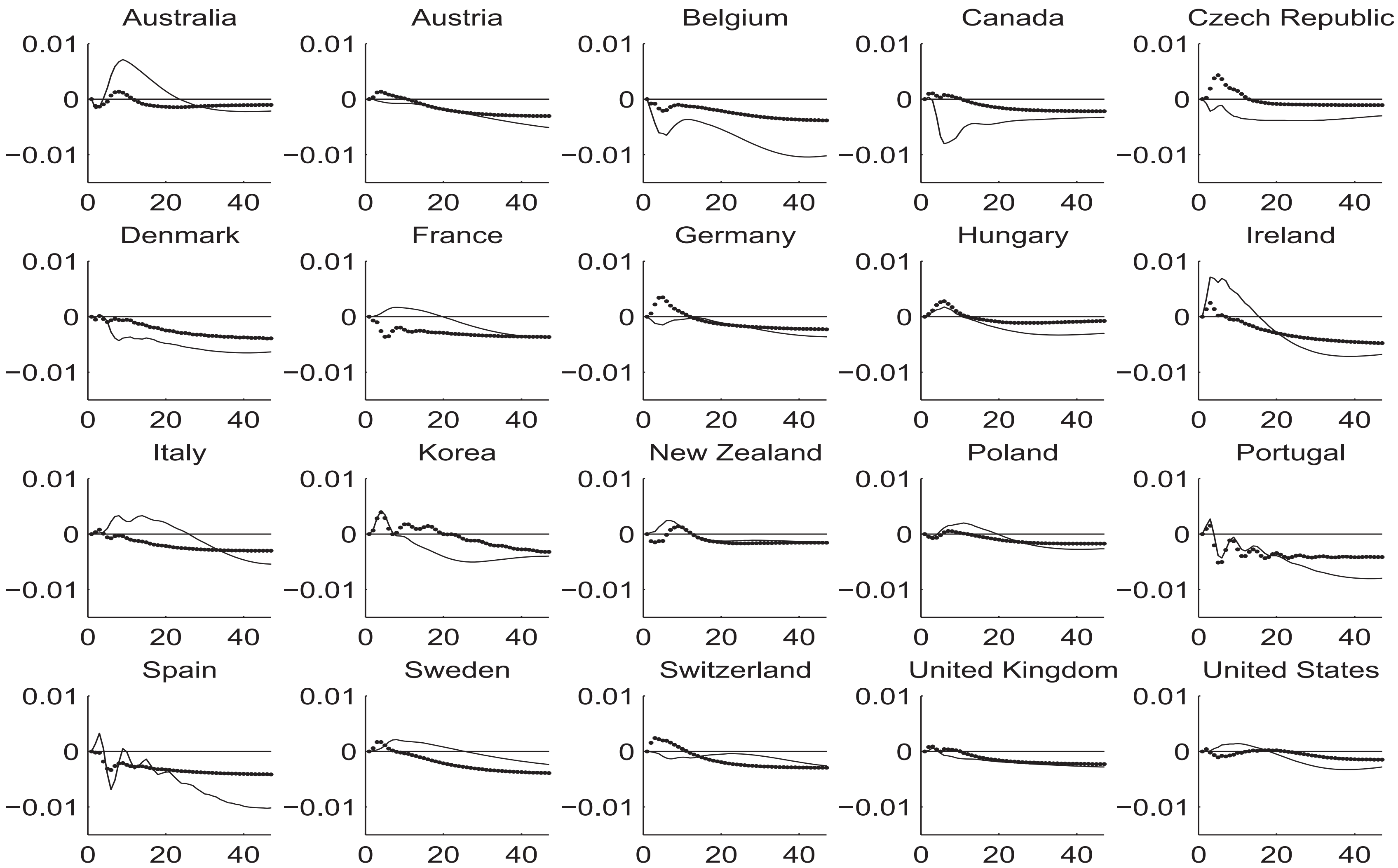
Note: The figure displays the price level responses to a 100 basis points monetary policy shock obtained from the the PCHVAR model in Equation (8). In each panel, the solid line depicts the point estimates of the impulse responses and the shaded area 90% bootstrap confidence bands.

Figure 7: Summary of Country VAR Model Output Responses to a Contractionary Monetary Policy Shock



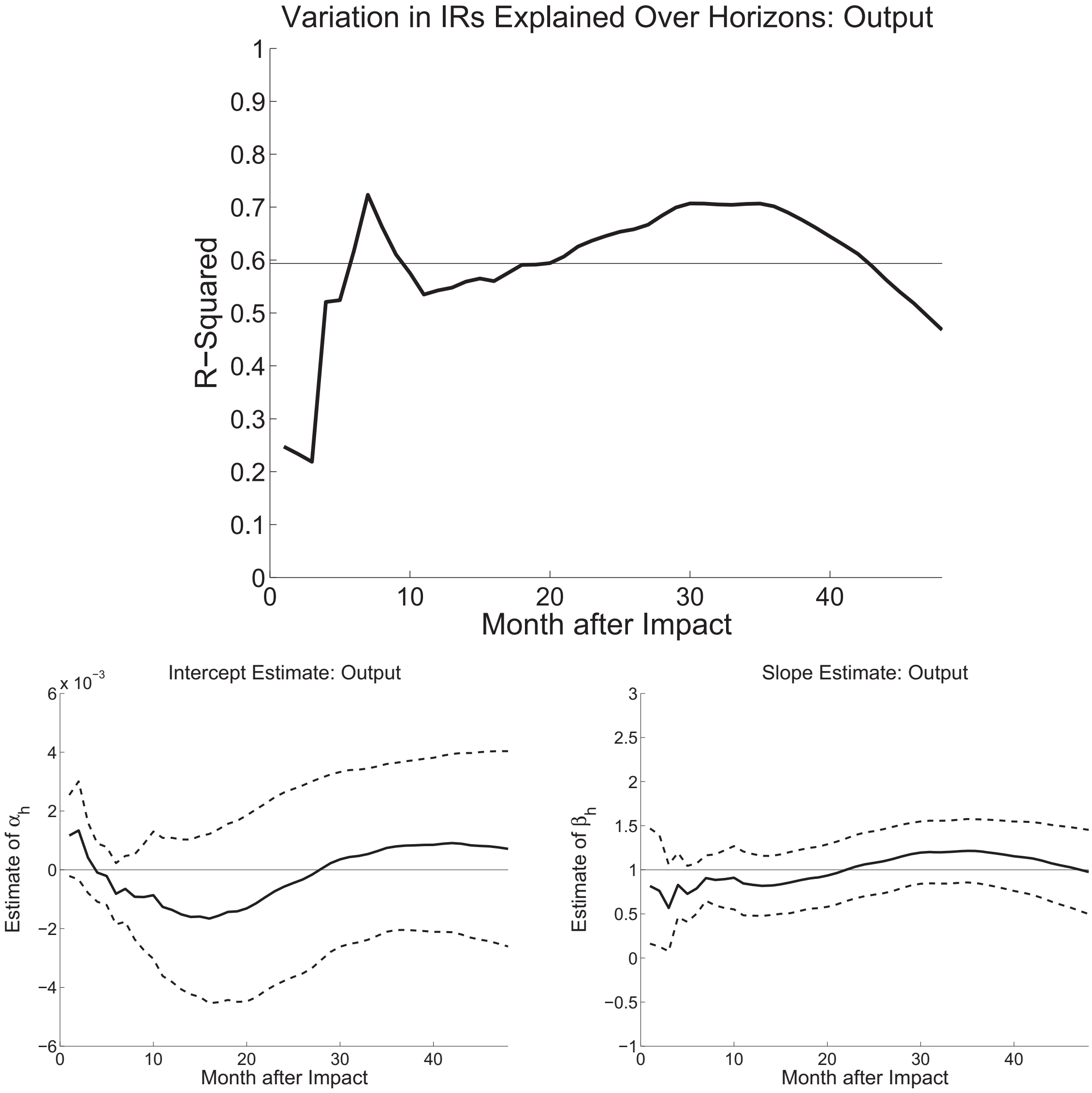
Note: The figure displays the output responses to a 100 basis points monetary policy shock obtained from the country VAR models in Equation (1) and the PCHVAR model in Equation (8). In each panel, the solid line depicts the response obtained from the country VAR model and the dotted line the response obtained from the PCHVAR model.

Figure 8: Summary of Country VAR Model Price Level Responses to a Contractionary Monetary Policy Shock



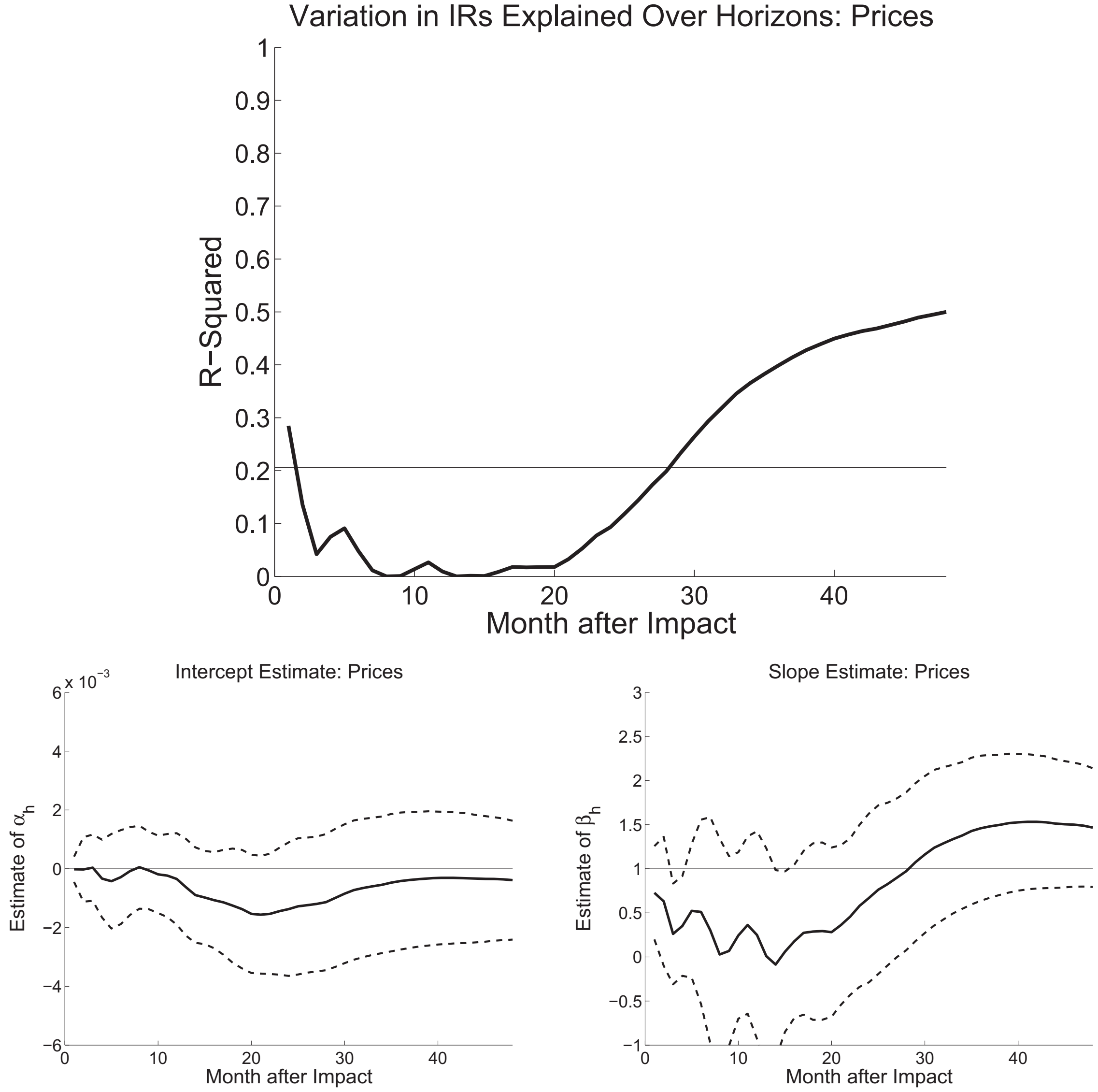
Note: The figure displays the price level responses to an exogenous 100 basis points monetary policy shock obtained from the country VAR models in Equation (1) and the PCHVAR model in Equation (8). In each panel, the solid line depicts the response obtained from the country VAR model and the dotted line the response obtained from the PCHVAR model.

Figure 9: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses of Output at Different Horizons



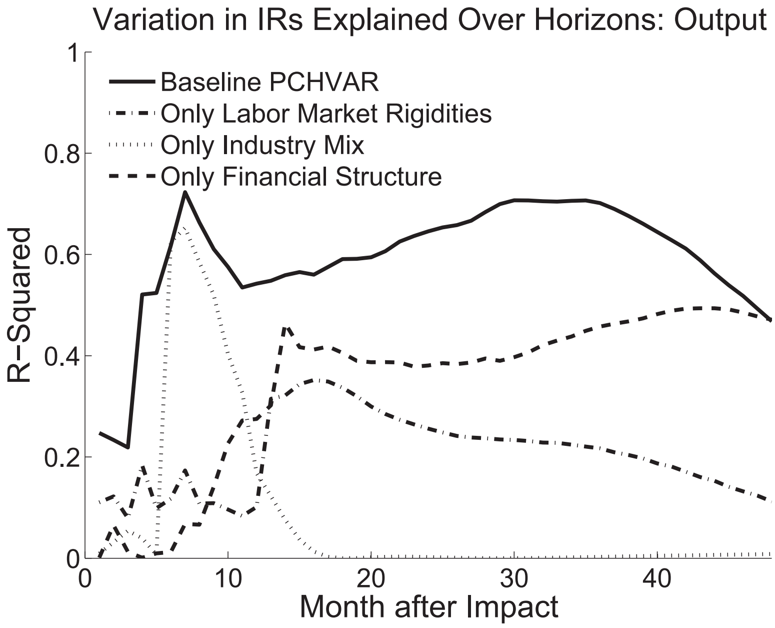
Note: The top panel displays the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) for output. The horizontal line represents the average of the R-squared over all response horizons. The bottom left-hand side panel depicts the evolution of the intercept estimate of Equation (10) (solid line) together with 90% confidence bands (dashed lines). The bottom right-hand side panel depicts the evolution of the slope of Equation (10) (solid line) together with 90% confidence bands (dashed lines).

Figure 10: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses of Prices at Different Horizons



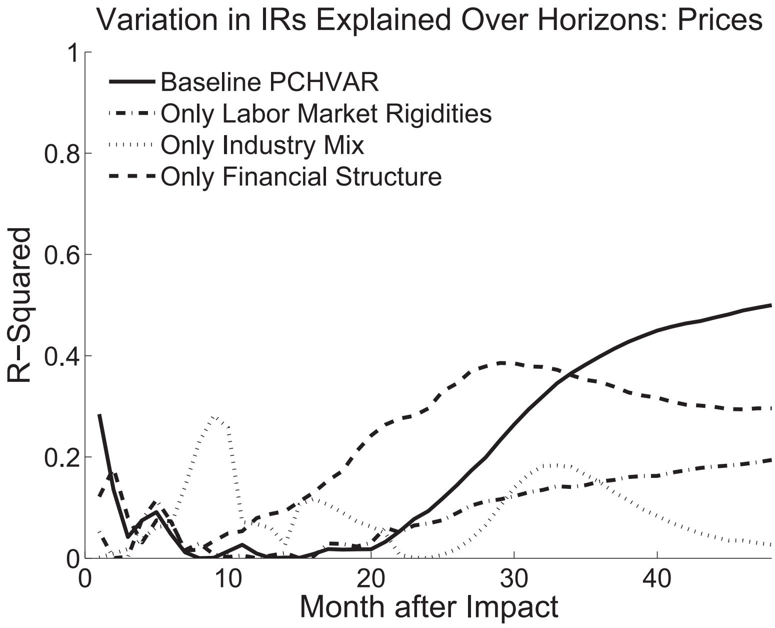
Note: The top panel displays the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) for prices. The horizontal line represents the average of the R-squared over all response horizons. The bottom left-hand side panel depicts the evolution of the intercept estimate of Equation (10) (solid line) together with 90% confidence bands (dashed lines). The bottom right-hand side panel depicts the evolution of the slope of Equation (10) (solid line) together with 90% confidence bands (dashed lines).

Figure 11: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses of Output at Different Horizons - Contributions



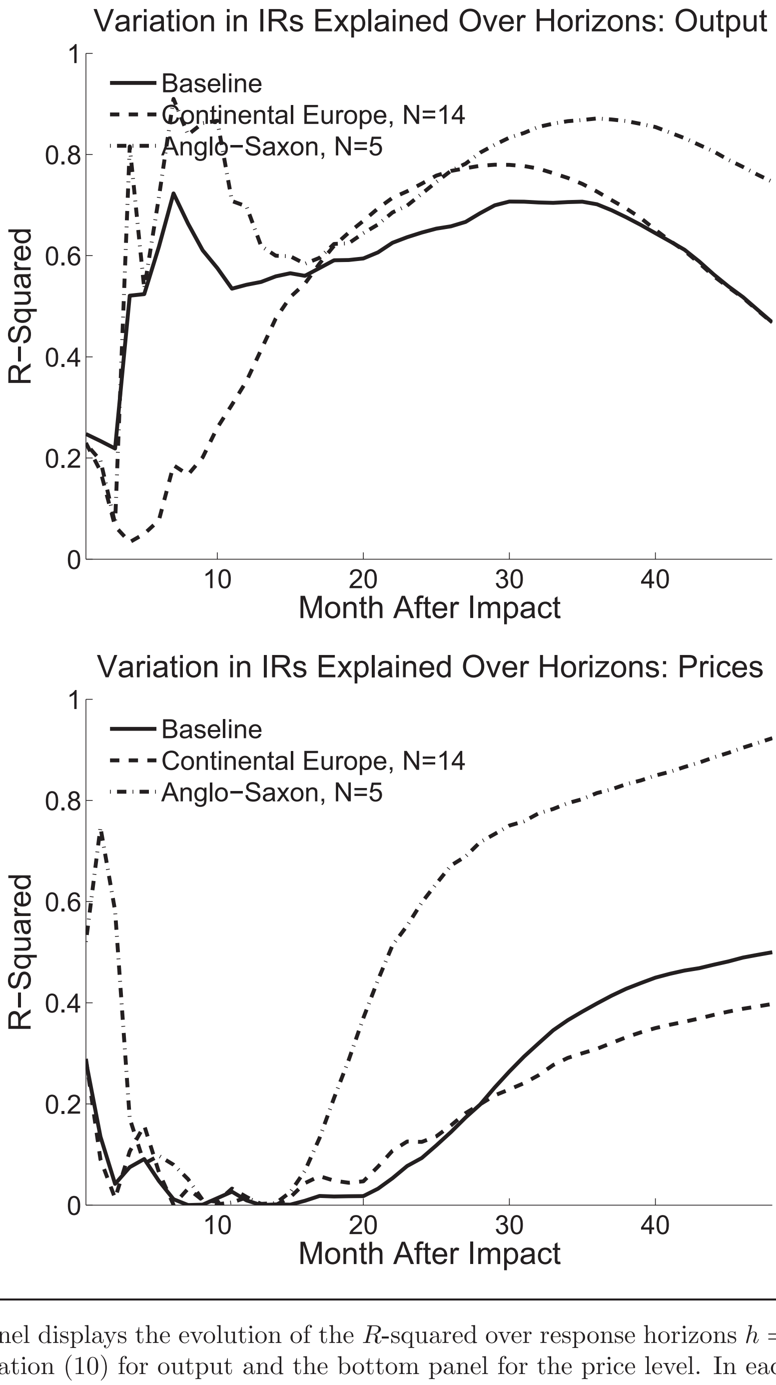
Note: The figure displays the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) for output. The solid line represents the results from the baseline specification with all thee structural characteristics considered jointly. The dash-dotted line represents the results from considering only financial structure, the dashed line those from considering only labor market rigidities, and the dotted line those from considering only industry mix. In order to reduce the impact of outlier observations on the R-squareds for the three latter cases, at each response horizon I drop the observation which is farthest away from the fit.

Figure 12: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses of Prices at Different Horizons - Contributions



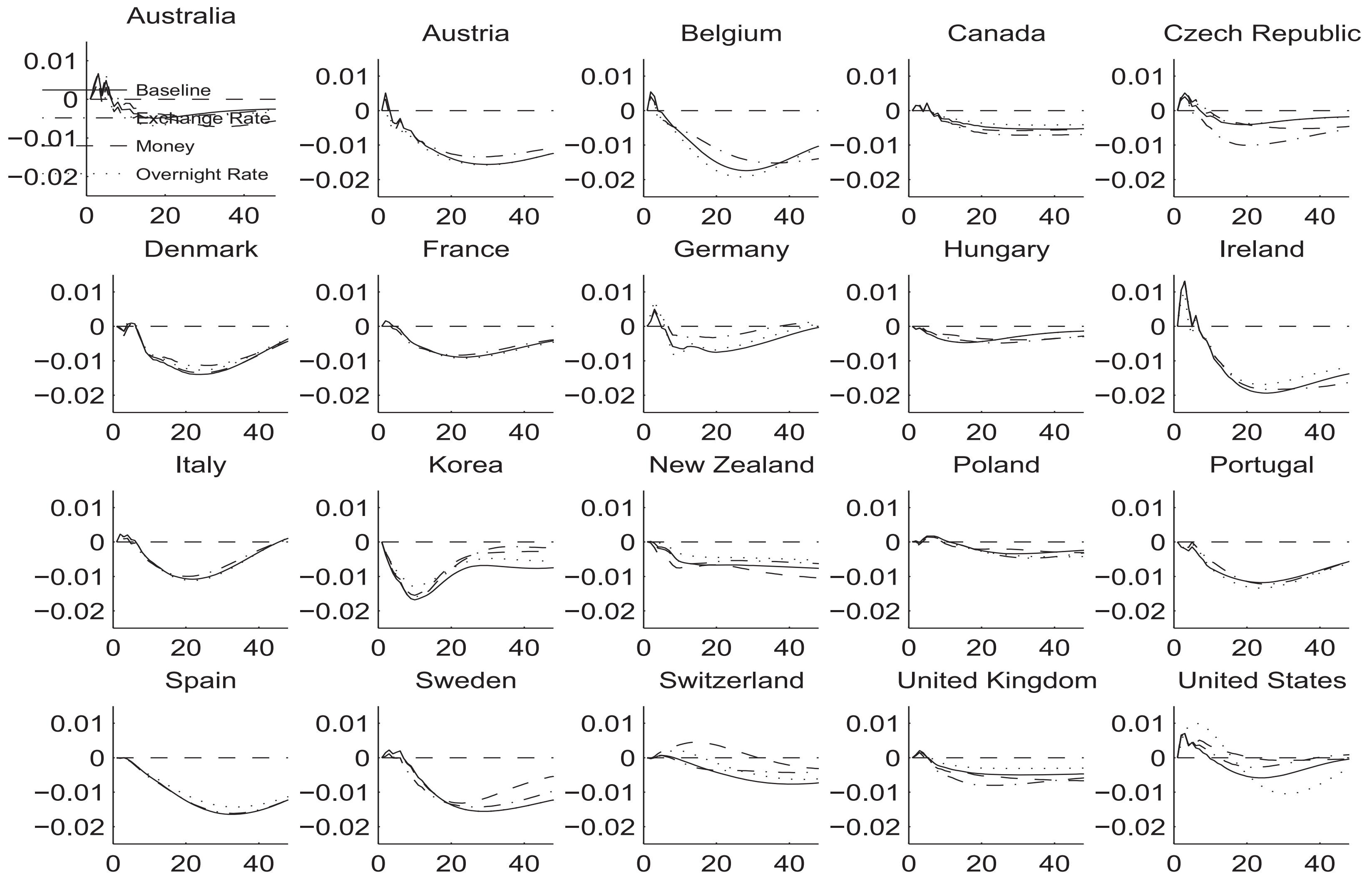
Note: The figure depicts the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) for prices. The solid line represents the results from the baseline specification with all thee structural characteristics considered jointly. The dash-dotted line represents the results from considering only financial structure, the dashed line those from considering only labor market rigidities, and the dotted line those from considering only industry mix. In order to reduce the impact of outlier observations on the R-squareds for the three latter cases, at each response horizon I drop the observation which is farthest away from the fit.

Figure 13: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses of at Different Horizons: Europe and Anglo-Saxon Countries



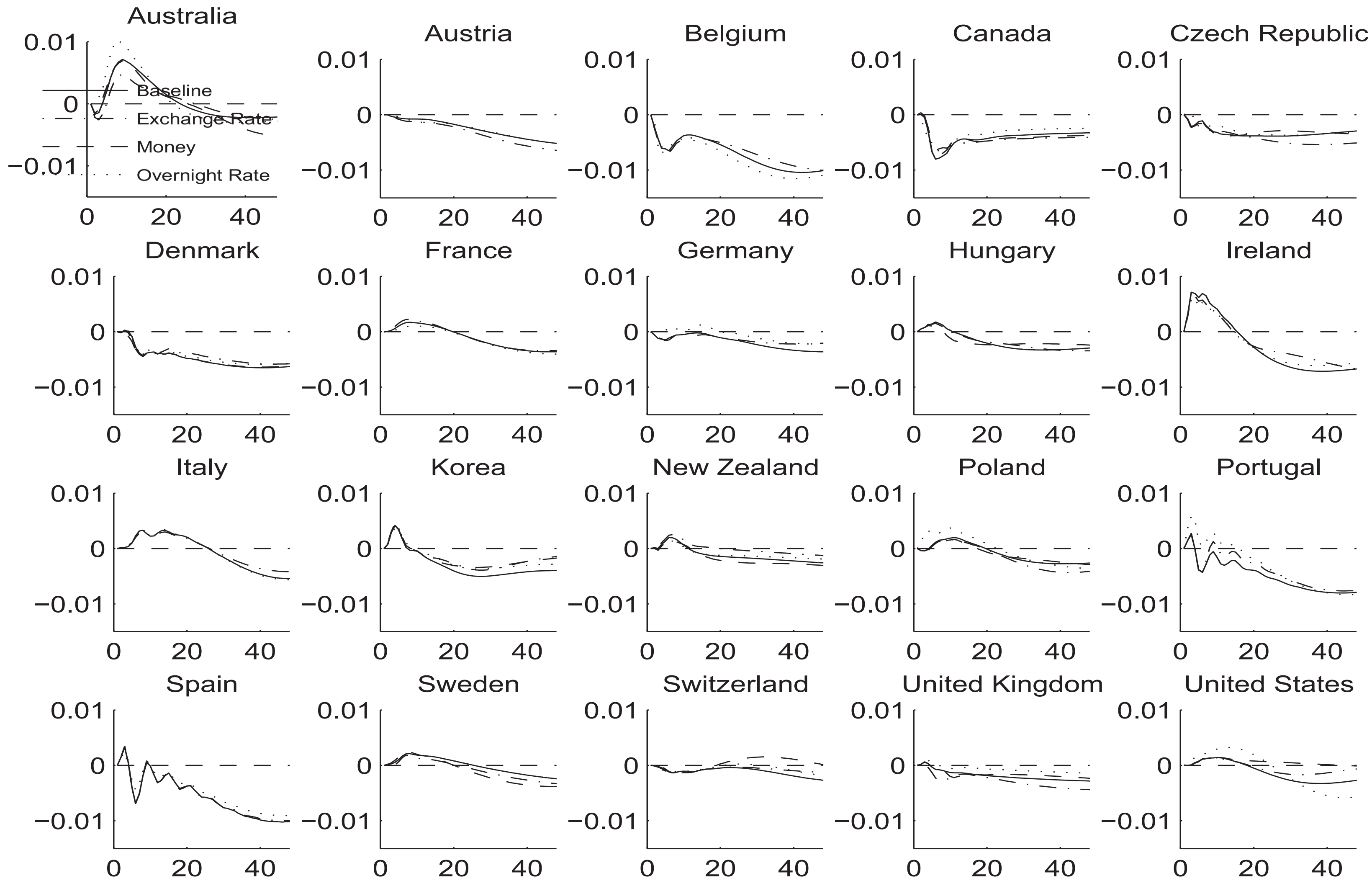
Note: The top panel displays the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) for output and the bottom panel for the price level. In each panel, the solid line represents the evolution of the R-squared for the baseline sample, the dashed line for the Continental European countries in the sample, and the dash-dotted line for the Anglo-Saxon countries in the sample.

Figure 14: Output Responses to a Monetary Policy Shock: Robustness Checks with Exchange Rates, Money and Overnight Rates



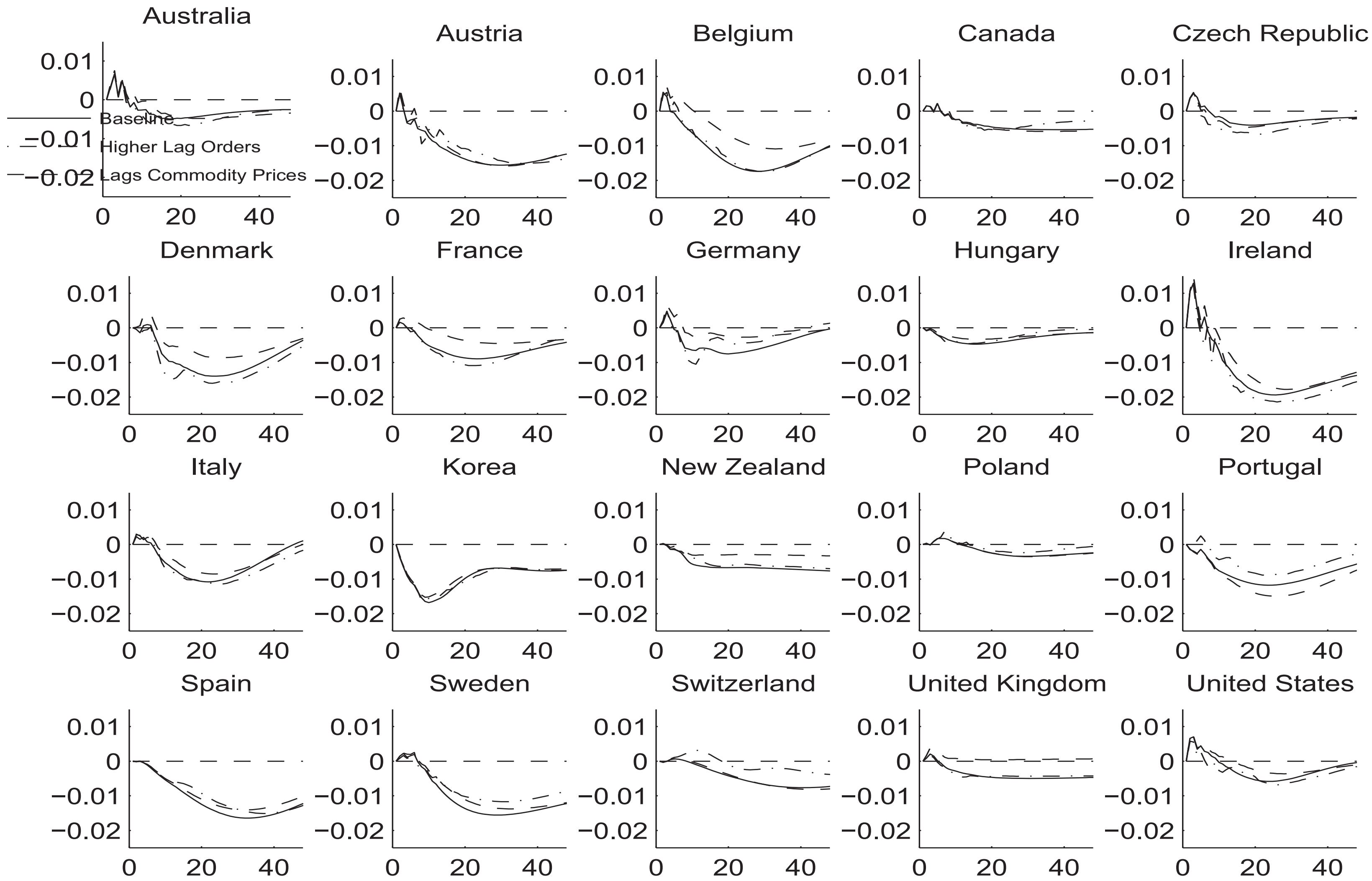
Note: The figure displays for all countries the output responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) (solid lines) together with the responses from country VAR models augmented by the nominal effective exchange rate (dash-dotted lines), money (dotted lines), and country VAR models in which the three-month money market rate is replaced by an overnight rate (dotted lines).

Figure 15: Price Level Responses to a Monetary Policy Shock: Robustness Checks with Exchange Rates, Money and Overnight Rates



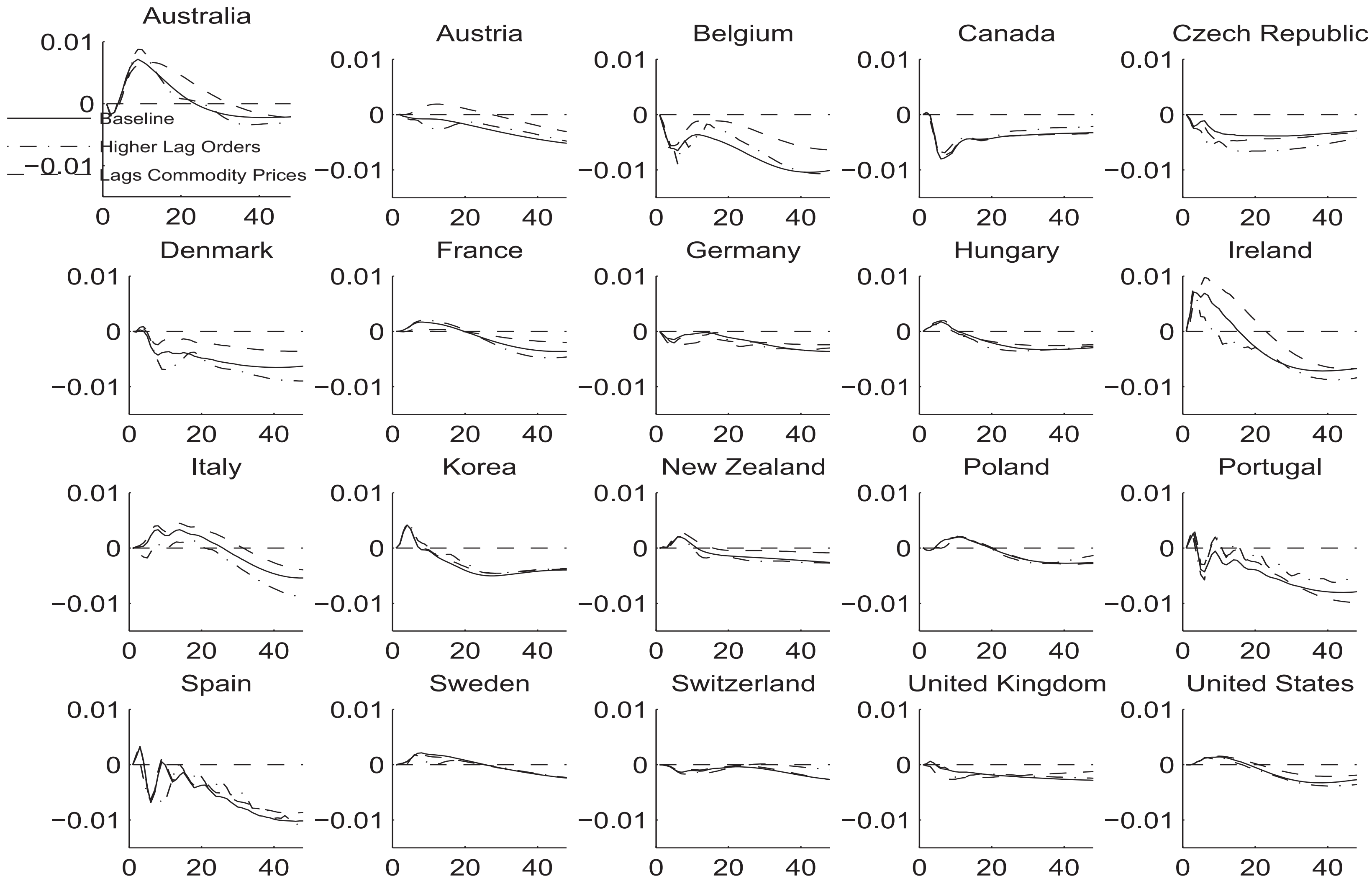
Note: The figure displays for all countries the price level responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) (solid lines) together with the responses from country VAR models augmented by the nominal effective exchange rate (dash-dotted lines), money (dotted lines), and country VAR models in which the three-month money market rate is replaced by an overnight rate (dotted lines).

Figure 16: Output Responses to a Monetary Policy Shock: Robustness Checks with Alternative Lag Orders



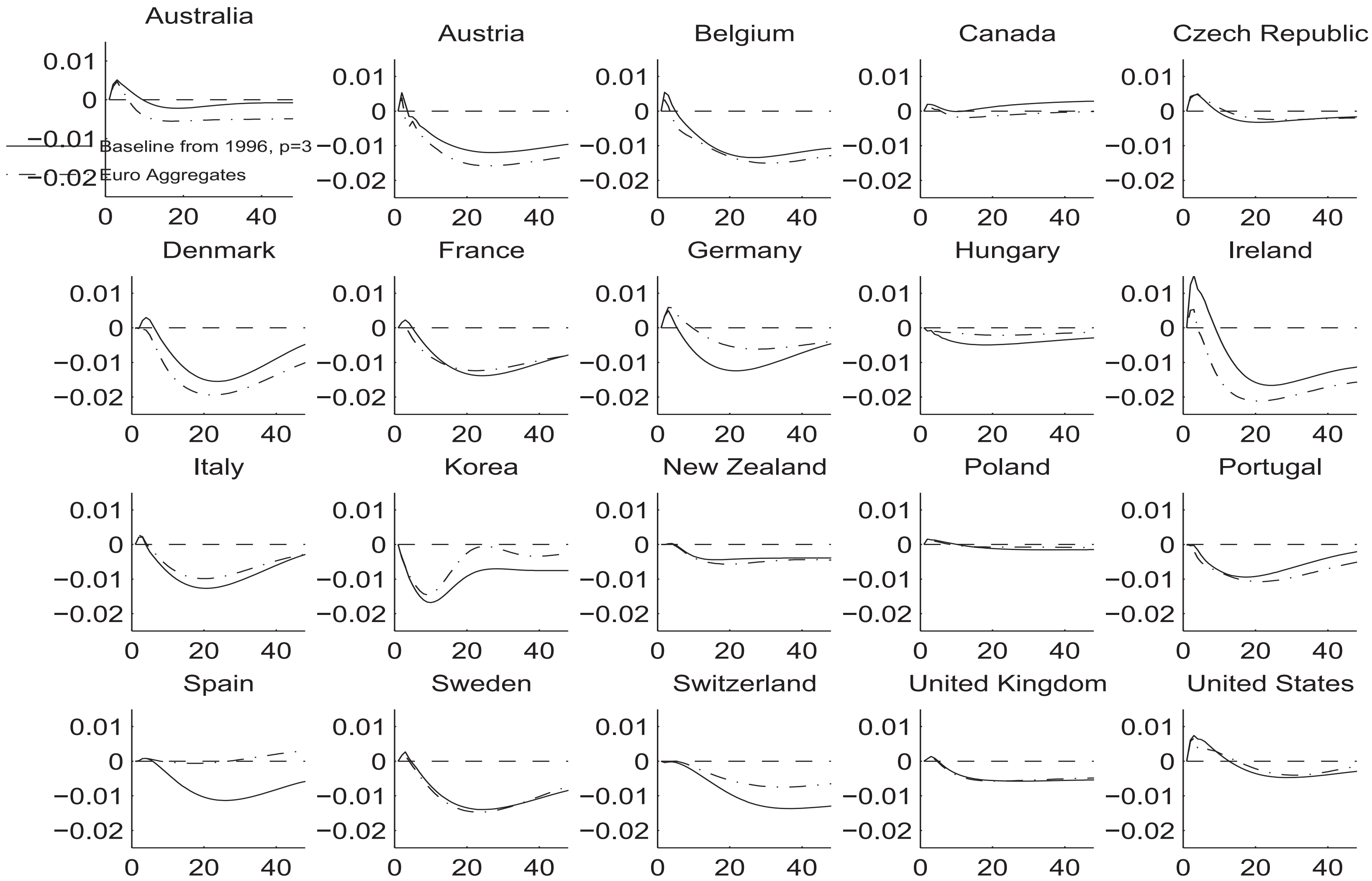
Note: The figure displays for all countries the output responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) (solid lines) together with the responses from the country VAR models with a lag order of nine (dash-dotted lines), and impulse responses with the lag order of the commodity price index increased to three (dashed lines).

Figure 17: Price Level Responses to a Monetary Policy Shock: Robustness Checks with Alternative Lag Orders



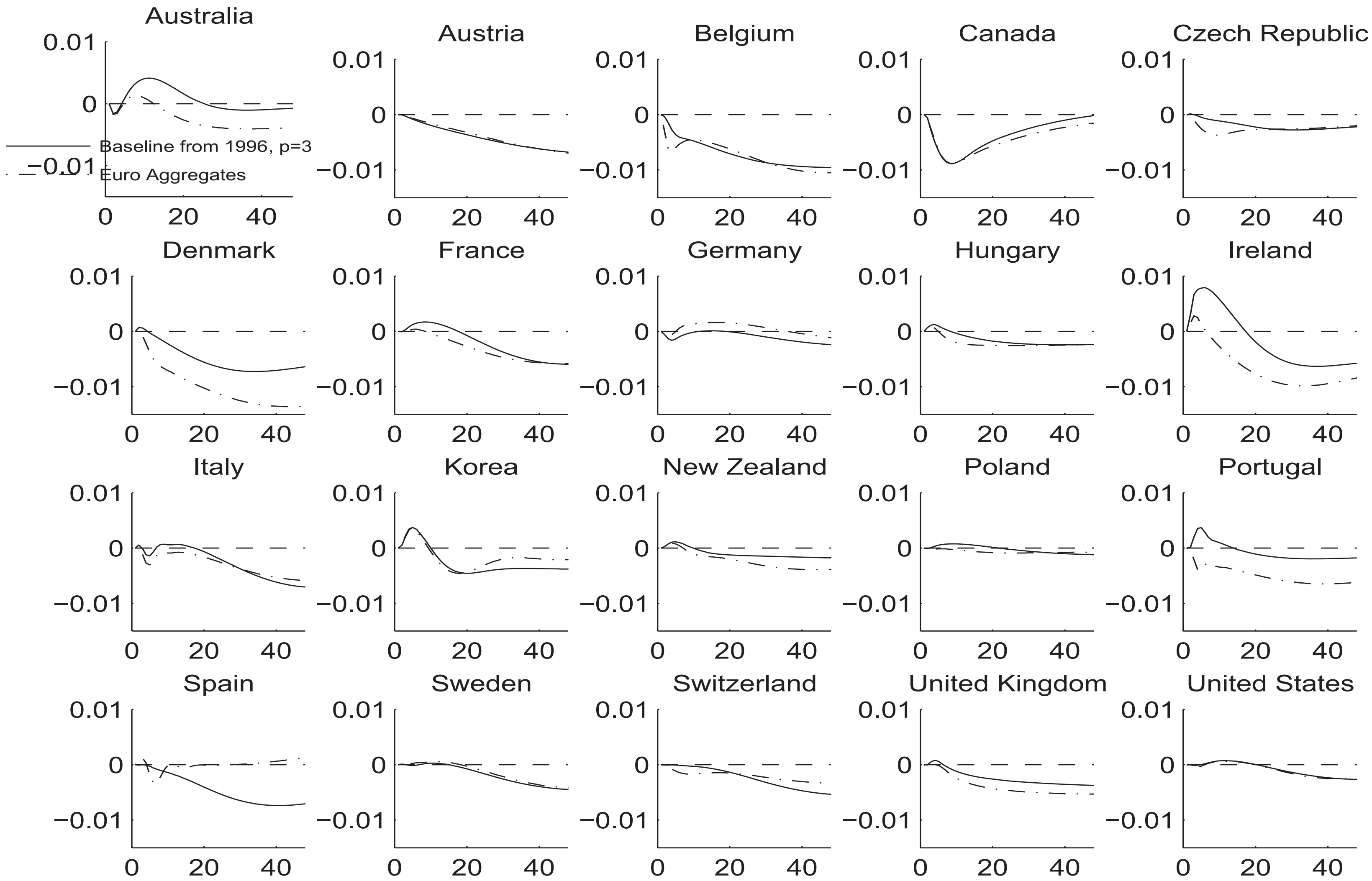
Note: The figure displays for all countries the price level responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) (solid lines) together with the responses from the country VAR models with a lag order of nine (dash-dotted lines), and impulse responses with the lag order of the commodity price index increased to three (dashed lines).

Figure 18: Output Responses to a Monetary Policy Shock: Robustness Checks with Euro Area Aggregates



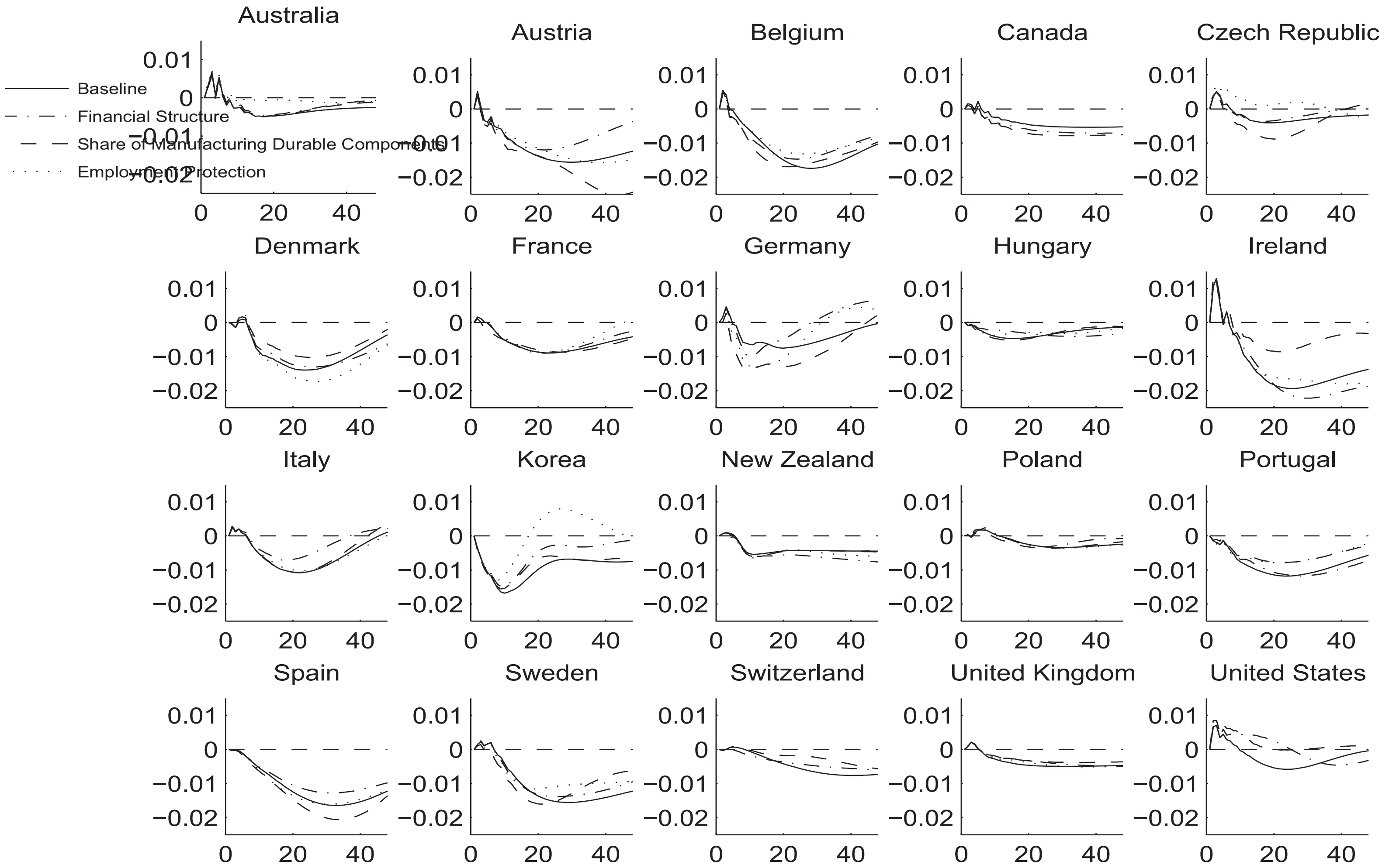
Note: The figure displays for all countries the output responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) estimated from 1996 (solid lines) together with the responses from the country VAR models augmented by the logarithms of euro area real GDP and the HICP (dash-dotted lines).

Figure 19: Price Level Responses to a Monetary Policy Shock: Robustness Checks with Euro Area Aggregates



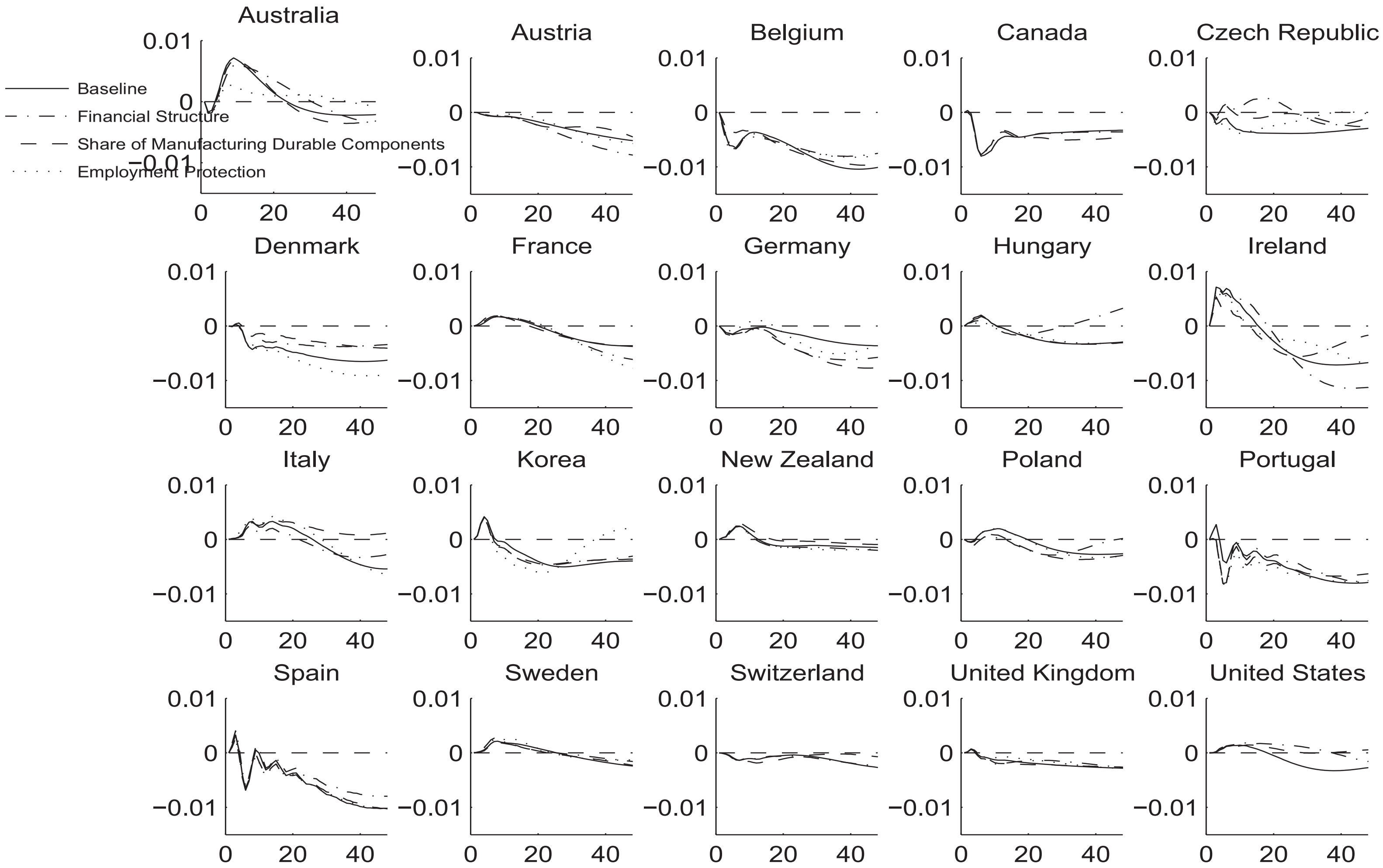
Note: The figure displays for all countries the price level responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) estimated from 1996 (solid lines) together with the responses from the country VAR models augmented by the logarithms of euro area real GDP and the HICP (dash-dotted lines).

Figure 20: Output Responses to a Monetary Policy Shock: Robustness Checks with Structural Characteristics



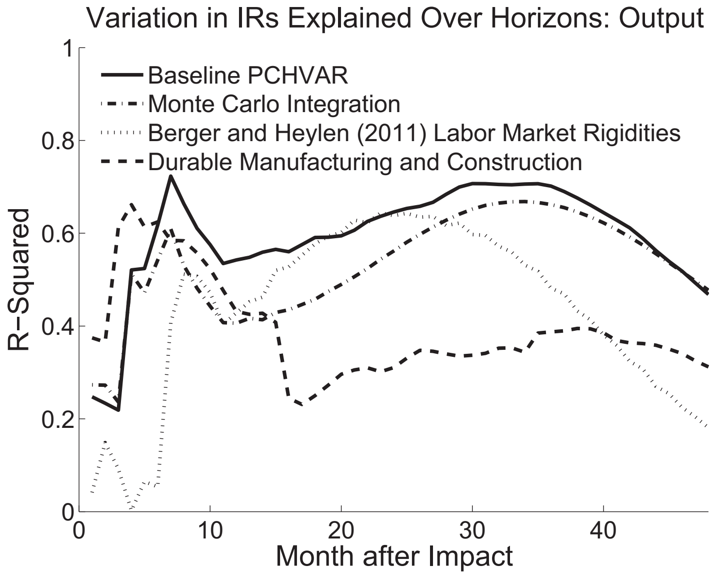
Note: The figure displays for all countries the output responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) (solid lines) together with the responses from the country VAR models augmented by the structural characteristics financial structure (dash-dotted lines), labor market rigidity (dashed lines) and industry mix (dashed lines) as additional endogenous variables one at a time. For the countries that do not feature any time-series variation in employment protection the corresponding impulse response can not be estimated and is therefore missing.

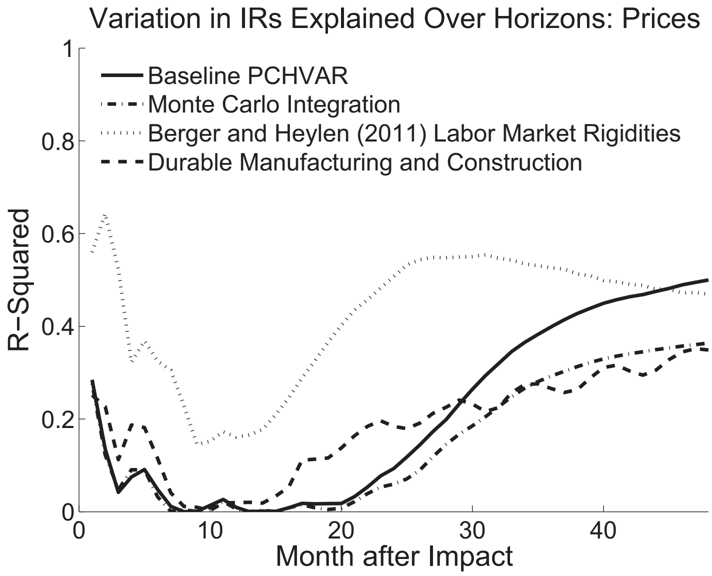
Figure 21: Price Level Responses to a Monetary Policy Shock: Robustness Checks with Structural Characteristics



Note: The figure displays for all countries the price level responses to a 100 basis points monetary policy shock obtained from the baseline country VAR models in Equation (1) (solid lines) together with the responses from the country VAR models augmented by the structural characteristics financial structure (dash-dotted lines), labor market rigidity (dashed lines) and industry mix (dashed lines) as additional endogenous variables one at a time. For the countries that do not feature any time-series variation in employment protection the corresponding impulse response can not be estimated and is therefore missing.

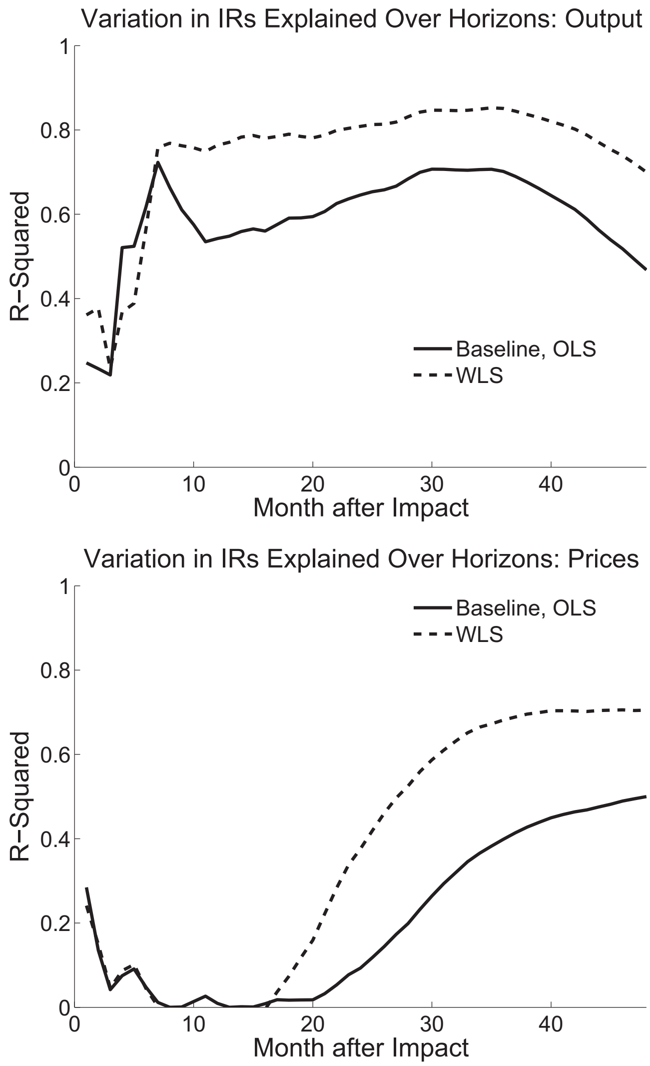
Figure 22: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses at Different Horizons: Robustness





Note: The figure displays the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) of the baseline country VAR and PCHVAR model specifications together with the corresponding results for the PCHVAR model impulse responses constructed using Monte Carlo integration along the lines of Koop et al. (1996) (dash-dotted lines). The dotted lines show the evolution of the R-squared when an alternative index of the intensity of labor market rigidities constructed on the basis of the data of Berger and Heylen (2011) is used in place of the OECD's Strictness of Employment Protection indicator.

Figure 23: The Relationship Between Country VAR Model and PCHVAR Model Impulse Responses at Different Horizons



Note: The upper panel displays the evolution of the R-squared over response horizons $h = 0, 1, 2, \ldots, H$ for the model in Equation (10) for output estimated by ordinary least squares (solid line) and weighted least squares (dashed line). In the weighted least squares case, the weight of each observation is the inverse of the sum of the standard errors of the country VAR model and the PCHVAR model impulse response estimates. The bottom panel displays the corresponding results for prices.

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