Sectoral Exchange Rate Pass-Through in the Euro Area*

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Abstract

We study exchange rate pass-through (ERPT) to euro area manufacturing import prices at the sectorally disaggregated level as well as for the manufacturing aggregate. Our estimation strategy is based on VAR models allowing for endogeneity of the explanatory variables. Endogeneity of the explanatory variables implies that often used single-equation OLS estimates of ERPT are inconsistent. The main findings are: large heterogeneity of ERPT across sectors and no evidence for instability of ERPT over time, contrasting some of the recent literature. When comparing our system based estimates with single-equation estimates substantial differences emerge, underlying the potential importance of using system methods. Furthermore, using VAR models and impulse response functions allows to study not only the extent but also the dynamics of ERPT.

JEL Classification: C50, F30, F40

Keywords: exchange rates, import prices, pass-through, euro area, sectoral disaggregation

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1 Introduction

Exchange rate pass-through (ERPT) is a focus of interest for policymakers and academics alike. The former are primarily interested in the extent and timing of ERPT as a key ingredient of their forecasting models of prices and of the trade balance. Along with academics, they are also interested in the role of ERPT in understanding the mechanisms of international price adjustment, e.g. in reconciling the observation that the relative stability of import prices does not reflect the high volatility of nominal exchange rates with economic theory. Evidence of 'disconnect' between exchange rates and prices would also imply a greater degree of insulation and thus greater effectiveness of monetary policy. Specifically, the main issue of interest is thus whether ERPT is *complete*, defined as a one-to-one response of import prices to exchange rate changes, and if so at what time horizon. A growing amount of empirical evidence suggests that pass-through is incomplete even in the long-run and much attention is spent on understanding the causes of incompleteness.[1]

We study ERPT to import prices at the sectoral level in the euro area, using data on 16 sub-sectors of manufacturing as well as the manufacturing aggregate. Due to the lack of import price data, however, we use import unit value indices (UVX) constructed dividing unit values of imports measured in national currency by import volumes in tons, which are available at a very disaggregated level from Eurostat's COMEXT database.[2] The data are briefly described in Section 3, while Appendix A gives details of the sources, weighting scheme and the aggregation steps. Our data set differs in several respects from data used in other studies, which stems from the fact that we use time-, sector- and partner specific weights in the aggregation schemes. The trading partners considered are, ordered by import weight (in total manufacturing): the USA, UK, Japan, Switzerland, Sweden, Hungary, the Czech Republic, Poland, South Korea, Denmark, Canada and Norway. The import weights are reported in Table 1 in Section 3. Though we consider only twelve partners, these countries cover almost 85% of euro area imports.[3]

[1] In this paper we define the exchange rate in terms of units of foreign currency per euro; as a consequence, our ERPT estimates are expected to have a negative sign and complete ERPT corresponds to a value of -1.

[2] Menon (1996) contains cautions against the use of price proxies - see also Alterman (1995) and Hooper and Richardson (1991).

[3] These import weights correspond to those used for the calculation of the ECB's official effective exchange rates and are based on average manufacturing imports in 1999-2001. The main part of the remaining 15.3% of imports is with China (9% of euro area imports in manufacturing), for which however data are particularly scarce and unreliable.

When discussing ERPT, it is important to qualify the prices of reference: while it is natural to expect incomplete ERPT to final consumer price indices, which contain a very high share of non-traded goods and services, a much debated question is what prevents exchange rates from feeding one-to-one into import prices even at longer horizons. The theoretical literature offers two main points of view: one strand points to nominal rigidities as the main cause of the non-responsiveness of import prices to exchange rate changes, while the other takes a more microeconomic point of view and ascribes incomplete ERPT to market and industry characteristics[4] and to the weight of non-traded inputs in the wholesale and consumer prices of imported goods.[5] The explanations stemming from an industrial organization point of view, while highlighting different mechanisms, hinge on the characteristics of the market considered. Specifically, ERPT is seen to depend, *inter alia*, upon the substitutability between foreign and domestic goods, the competitive structure in the industry (again domestically and abroad), barriers to trade, import penetration and the relative market size. If exporters set prices in their own currency, labelled producer currency pricing (PCP), then ERPT is 1 in the short run.[6] If exporters set prices in the currency of the destination country, labelled as local currency pricing (LCP), then ERPT is 0 in the short run. LCP is a particular form of 'pricing to market' (PTM), which describes the behavior of exporting firms who change their mark-ups depending on the destination markets, offsetting exchange rate movements either completely (LCP) or only to some degree, e.g. to maintain market shares. In fact, in a dynamic setting the response pattern may be more complicated than suggested by the clear-cut boundary cases of PCP and LCP, as firms may have to adjust prices in either case over time in response to exchange rate fluctuations. It is also important to note here that a large part of the industrial organization–trade type literature on ERPT is based on static models and often places (usually implicitly) exogeneity restrictions on foreign variables and the nominal exchange rate. Therefore, in the empirical assessment of the relationships highlighted by this type of literature, single-equation OLS estimation strategies are usually pursued. If the explanatory variables are, however, endogenous, only a systems approach (as pursued in this paper) or instrumental variable estimation allows for consistent estimation of

[4] Key contributions in this spirit are Krugman (1987), Dornbusch (1987) or Baldwin (1988).

[5] See e.g Burstein, Eichenbaum, and Rebelo (2005), Burstein, Neves, and Rebelo (2003), or Goldberg and Verboven (2005).

[6] With our definition of nominal exchange rates (in terms of foreign currency per unit of domestic currency), this corresponds in our paper to an ERPT of -1.

ERPT. In Section 4.1 we compare the results from our systems approach with single-equation OLS findings. Given that our exogeneity tests clearly reject the null hypothesis of exogeneity of the explanatory variables, single-equation OLS estimation on our data set is bound to produce inconsistent estimates. Specifically, in Tables 8 and 9 in Section 4.1 we gauge the impact of the absence of exogeneity on the single-equation OLS estimates by comparing them with our VAR-based estimates.

More recently, new open economy macroeconomics (NOEM) models, e.g. Corsetti, Dedola, and Leduc (2005), have incorporated more micro-founded explanations in dynamic stochastic general equilibrium models allowing for different degrees of nominal rigidities. These models are typically set in a two-country framework and endogenously determine the optimal degree of ERPT as a function of e.g. the share of distribution costs in the final price of imported goods, the degree of substitutability between goods, and other factors affecting steady-state mark-ups. Typically, these models trade in details of market structure characteristics for an inter-temporal optimization model and thus they are more related to price aggregates than the industrial organization and trade theory type of models. However, due to its dynamic character, this literature is more amenable to be investigated with dynamic econometric models like vector autoregressive (VAR) models.

In our paper we follow a somewhat hybrid approach: we use sectorally disaggregated data to study whether pass-through differs across 'industries'[7] and we use VAR models that contain additional explanatory variables like commodity prices or some measure of the output gap to control for economy-wide supply- and demand-side characteristics to take into account (at least to some extent) the fact that prices and exchange rates are jointly determined in a general equilibrium setting.

Another issue that has spurred quite some discussion in policy and academic circles is whether ERPT has declined over time, especially since the early 1980's. The literature proposes two types of explanations: The first attributes declines in pass-through to the fact that the currencies of countries with stable monetary policy and low inflation are more likely to exhibit low ERPT.[8] The second explanation (proposed by e.g. Campa and Goldberg, 2005) is based on sectorally different ERPTs. In the presence of different degrees of ERPT across

[7]We use the term 'industries' in quotation marks because the sectors we study are too aggregated to be really considered as industries.

[8]See Taylor (2000) and Engel (2005) for theoretical analyses of this argument and Gagnon and Ihrig (2004) for an empirical assessment.

sectors, a switch in the import composition towards low pass-through sectors may lead to a decline in observed aggregate pass-through. Our sectoral analysis provides valuable input for this line of research by providing a quantification of the inter-sectoral variability of pass-through.

We investigate the issue of declining ERPT over time by conducting stability analysis for our VAR models. Our results, briefly discussed in Section 4 and available in detail upon request do not indicate that structural change occurs at the sectoral level. This finding is in some contrast to the conclusions of Campa and Goldberg (2005), who find evidence in favor of this hypothesis for OECD countries. However, their findings are based on single-equation OLS results, which may consequently suffer from endogeneity bias.

Investigating pass-through to import prices is to be seen merely as the first step of an analysis of the impact of exchange rate changes on domestic prices. Subsequently it is of particular interest to study in more detail the pricing chain from import to wholesale to consumer prices (as done by e.g. Hahn (2003) for euro area price aggregates). In our paper we do not follow this route but focus in detail (with different specifications and identifying assumptions) on the pass-through to import prices.

We find sizeable differences in ERPT across sub-sectors of manufacturing. Contemporaneous ERPT varies from 0.16 (not significantly different from 0) for radio and television equipment to -0.81 (not significantly different from -1) for wearing apparel.[9] For twelve of sixteen sectors contemporaneous ERPT is not significantly different from 0 and for 2 it is not significantly different from -1. The long-run (to be precise after two years) ERPT varies from -0.12 (not significantly different from 0) for food products and beverages to -1.57 (not significantly different from -1) for fabricated metals. Long-run ERPT is not significantly different from 0 in six sectors and not significantly different from -1 in ten sectors. The uncertainty around the estimates also varies considerably across sectors, see the detailed discussions in Section 4. For both aggregate measures long-run ERPT is not significantly different from -1, however short-run ERPT differs between the two measures.

The paper is organized as follows: In Section 2 we provide a brief overview of existing work on sectoral ERPT to import prices, while Section 3 discusses the data and the modelling framework. The empirical results are presented in Section 4. Section 5 briefly summarizes and

[9] Two sectors produce implausibly large contemporaneous ERPT estimates. These are the MIN (+2.00) and the PETR (-1.50) sectors.

concludes. Two appendices follow the main text: Appendix A gives further details concerning the data and the construction of the variables and Appendix B contains additional empirical results.

2 Previous Studies on Sectoral ERPT

We are not the first to conduct an empirical study of ERPT at the sectoral level, although we are, to the best of our knowledge, the first to do it for the euro area as a whole and to depart from using a variant of a simple single-equation regression. Previous studies based their empirical analyses on various versions of the equation discussed by Goldberg and Knetter (1997) that relates the price of imports (or exports, depending on the point of view) to the nominal exchange rate, a primary explanatory variable which is a measure of domestic prices, and other variables labelled 'demand shifters' (sometimes proxied by GDP growth). This type of equation is usually estimated in a single equation framework.

The degree of sophistication of such empirical approaches varies considerably. In order to facilitate the comparison of our results with those of other studies, we briefly present below the equations estimated in some papers that study sectoral pass-through.

Knetter (1993) and Yang (1997) are two rather early studies which look at the problem from the point of view of exporting firms. Knetter (1993) uses a two-way fixed effects model to study the pricing to market behavior of US, UK, German and Japanese exporting firms using unit values of exports for rather disaggregated (seven digits) industries and finds more variation in the degree of ERPT across industries than across countries. He simply regresses the growth rates of prices on those of exchange rates, while exporters' marginal costs are meant to be captured by the time effects in his panel regression.

Yang (1997) also looks at ERPT from the point of view of the exporter and studies US manufacturers across industrial sectors (which largely overlap with ours) and finds evidence of PTM behavior, i.e incomplete ERPT, with largely varying degrees across industries. He estimates the following equation for each sector:

$$\Delta p_t^m = \alpha \Delta e_t + \rho \Delta p_{t-1}^m + \beta \Delta p_t^d + \varepsilon_t \quad (1)$$

where p_t^m are import prices, e_t is the exchange rate and p_t^d are domestic prices.

More recently, Campa and Goldberg (2005) study ERPT to import prices in 23 OECD countries both at the aggregate and at a broadly disaggregated level. They study food, manufacturing, energy, raw materials and non-manufacturing imports and find evidence of partial ERPT in the short run, in particular in the food and manufacturing sectors. They estimate the following regression for each country and sector by OLS:

$$\Delta p_t^m = \mu + \sum_{i=0}^{4} \alpha_i \Delta e_{t-i} + \sum_{i=0}^{4} \beta_i \Delta c_{t-i}^* + \gamma \Delta gdp_t + \varepsilon_t \tag{2}$$

where c_t^* is a measure of foreign costs, which they derive implicitly from the ratio of the real to nominal exchange rate. The authors also investigate the stability over time of their estimated relationship and find little evidence of instabilities. Consequently they conclude that the discussed decline in ERPT observed in OECD countries since the 1980-90s is more due to a shift in the composition of imports away from high ERPT sectors like energy and into lower ERPT sectors like manufacturing and food.

Otani, Shiratsuka, and Shirota (2003) estimate an equation containing similar regressors, but include a lagged dependent variable:

$$\Delta p_t^m = \mu + \rho \Delta p_{t-1}^m + \alpha \Delta e_t + \beta \Delta c_t^* + \gamma \Delta iip_t + \varepsilon_t \tag{3}$$

where c_t^* is constructed implicitly as in Campa and Goldberg (2005) based on the real unit labor cost (ULC) and nominal effective exchange rates published by the IMF (using the inverse definition to the one used in this paper, i.e. in terms of units of foreign currency) and iip_t is the index of industrial production. They produce contrasting evidence on Japanese import prices to that proposed by Campa and Goldberg (2005), claiming that the decline they observe in ERPT does not come from a shifting composition of imports, but from a decline of ERPT at the product category level. They analyze food, raw materials, fuels, chemicals, textiles, metals and machinery and find some long-run estimates larger than 1, e.g. for food and raw materials.

Marazzi et al. (2005) estimate:

$$\Delta p_t^m = \mu + \sum_{i=0}^{K} \alpha_i \Delta (e_{t-i} + c_{t-i}^*) + \delta \Delta p_t^{\text{commodities}} + \varepsilon_t \tag{4}$$

where p_t^m are import prices of goods excluding oil, computers and semiconductors. They use foreign headline CPI as their measure of foreign cost, c_t^*, aggregated in the same way as the nominal effective exchange rate, via rolling non-oil import weights. In their robustness checks, they also add US PPI as a measure of domestic competitors' prices. They are primarily interested in the stability of ERPT over time and report that their estimates exhibit a decline since the late 80's, especially for consumer and capital goods and for automotive products. This is more in line with the results of Otani, Shiratsuka, and Shirota (2003) than with those of Campa and Goldberg (2005).

Campa, Goldberg, and González-Mínguez (2005) use the same methodology and estimate the same equation as Campa and Goldberg (2005) but use data for several euro area countries, both at the aggregate level and for nine sectors.[10] They find evidence of incomplete ERPT both in the short- and long-run across all sectors. They present also some simple averages of ERPT across euro area countries, but do not analyze euro area aggregates directly.

As mentioned above, a potentially large drawback of the empirical literature surveyed here is that single-equation OLS estimates of the coefficients are inconsistent if any explanatory variable is endogenous.

In fact, while our approach is similar to that of the existing literature as regards the choice of variables and its theoretical underpinnings, our choice of the VAR modelling framework not only allows us to study the dynamics of ERPT in more detail, but also equips us with the necessary tools to investigate empirically the question concerning exogeneity of the exchange rate and of domestic and foreign producer prices, which we choose as proxies for the price of substitutes and for foreign costs respectively. As mentioned above, the results of our comparison of findings obtained with VAR models and by single-equation OLS estimation are contained in Section 4.1. We do find quite some differences between the single-equation and VAR results, with no systematic patterns in the differences emerging across sectors.

From a more agnostic point of view, our results based on well-specified linear dynamic models can be seen as generators of stylized facts concerning the behavior of import prices (or to be precise unit value indices) with respect to various shocks. One of these shocks, which we focus on here, is a shock to the nominal effective exchange rate. However, there are other interesting effects that can be studied within the VAR framework. These include the effects of oil price or foreign cost/price shocks on import prices or domestic producer prices. In subsequent work we will extend our analysis to estimate models that impose more structure than in the present paper. In this study we focus on Cholesky decompositions to identify the 'structural' shocks and perform robustness analysis with respect to the shock ordering as well

[10]The sectors are: food and live animals, beverages and tobacco, inedible crude materials, mineral fuels, oils fats and waxes, chemical products, basic manufactures, machines and transport equipment and other manufactured goods.

as with respect to modelling the effects of additional explanatory variables, as discussed in Section 3.

3 Data and Modelling Framework

As discussed above, our choice of variables is quite standard in the empirical literature on ERPT to import prices: in most cases, empirical specifications include measures of foreign cost, prices of domestic substitutes and 'demand shifters'. We proxy foreign costs by the sector-specific foreign producer price index (PPI*) and the price of domestic substitutes by the domestic sector-specific producer price index (PPI). Of course, producer price indices comprise both costs and mark-ups and therefore it is clear that PPI* is a rough measure for cost. In contrast to the studies mentioned in Section 2, however, we include *all* these variables, while e.g. Yang (1997) omits a measure of exporters' costs and the other studies surveyed do not include a measure of the price of domestic substitutes. Furthermore we include several commodity price indices and the output gap in the modelling procedure, see Figure 4 and Table 3.

As mentioned in the introduction, the euro area trading partners considered in this study cover almost 85% of extra-euro area manufacturing imports, with weights as indicated in Table 1.

While the import data classification in COMEXT is based on SITC 3 product categories, we construct sector aggregates that match, to the extent possible, the ISIC 3.1 industry classification, in which disaggregated producer price index (PPI) data are available for euro area members as well as for the twelve trading partners. Table 2 lists the sub-sectors of manufacturing analyzed and the matching between ISIC 3.1 and SITC 3.

We form 'effective' foreign PPI and nominal exchange rate indices for each euro area country by geometric weighted averages across the twelve partners for each euro area member[11] and then we form the euro area aggregate by a 'second-level' geometric weighted average (see Appendix A for details on the weighting schemes).

Notwithstanding the drawback of using unit value indices instead of import prices, our data set offers, with respect to the data used in most of the empirical literature, the advantage of following the same aggregation scheme for both sectoral foreign PPI and the effective

[11] Belgium and Luxembourg are considered as one country since they are considered together in the trade data.

Country	Weight
UK	22.0%
USA	21.4%
Japan	10.9%
Switzerland	7.8%
Sweden	5.6%
Hungary	3.3%
Czech Republic	3.2%
Poland	3.0%
South Korea	2.8%
Denmark	2.4%
Canada	1.2%
Norway	1.1%
Total	84.7%

Table 1: Weight of each trading partner considered in extra-euro area manufacturing imports, average over 1999-2001.

nominal exchange rate. Also in contrast to the literature to the best of our knowledge, we construct time-, sector- and partner-specific weights for each euro area country based on volumes, rather than values of imports, to avoid the pitfall of currency-dependent weights. Furthermore, we construct time-varying monthly weights to incorporate changes over time in the country composition of euro area imports.[12] All these steps should allow for a 'cleaner' study of ERPT at the sectorally disaggregated level. Given the noisiness of in particular the import data, we adjust the variables for outliers and seasonality.

For completeness and comparison we also report estimation results for the aggregate data used for the construction of the official euro real PPI-deflated effective exchange rate. The latter is constructed using simple import weights and double export weights based on manufacturing trade values in 1999-2001.[13] Since export weights are also included in the calculation, the official effective exchange rate is not optimal for studying import price pass-through and this difference in the construction of the data may well be responsible for the observed dif-

[12]We also use a changing composition for the 'synthetic' euro area itself, given limited availability of data for some variables and some countries; e.g. Austrian PPIs are only available from 1996.

[13]For details on the construction of the weights used for the computation of the official effective exchange rates of the euro published by the ECB, see Buldorini, Makrydakis, and Thimann (2002).

Sector	ISIC Rev. 3	Description	SITC 3 (approximate)
CHEM	24	Chemicals and chemical products	59
COMP	30	Office machinery and computers	75
ELEC	31	Electrical machinery and apparatus n.e.c.	77
FABMET	28	Fabricated metal products, except machinery and equipment	69 - 699
FOODBEV	15	Food products and beverages	0, 00, 11 (subtract 00 from 0)
MACH	29	Machinery and equipment n.e.c.	71, 72, 73, 74
METALS	27	Basic metals	67 68 699
MIN	26	Other non-metallic mineral products	66
MOTOR	34	Motor vehicles, trailers and semi-trailers	78
PAPER	21	Pulp, paper and paper products	64
PETR	23	Coke, refined petroleum products and nuclear fuel	32, 33 (missing nuclear)
PLAST	25	Rubber and plastic products	62 58
PRECINSTR			
RADIOTV	32	Radio, tv, communication equipment/apparatus	76
TEXT	17	Textiles	65
WEAR	18	Wearing apparel; dressing and dyeing of fur	84

Table 2: List of sectors studied, with corresponding match between ISIC 3.1 and SITC 3 classifications.

ferences in short-run ERPT for the official and our exactly aggregated series. Furthermore, using monthly weights our aggregate precisely reflects the actual country composition of euro area imports.

In addition to domestic and foreign PPI, we also consider additional explanatory variables intended to capture demand side characteristics: we conduct specification searches using the output gap, GDP growth[14], M3 growth and the 3-month nominal interest rates. The output gap is defined in this paper as the deviation from the Hodrick-Prescott (HP) trend of real GDP.[15] Among these variables the output gap has the highest explanatory power.

As mentioned above we also include several commodity price indices that appear significant in explaining the import UVX as well as other variables in the VAR. As expected such variables increase explanatory power in the equations for the producer price indices substantially. These price variables are significant in many cases in addition to the oil price (compare again Table 3). Details on all these additional explanatory variables are given in Table 11 in Appendix A.

Corsetti, Dedola, and Leduc (2005) provide an assessment of the bias arising in empirical ERPT equations from omitting relevant variables such as marginal costs, or from proxying them with large error. In order to quantify this bias, they simulate data from their calibrated structural model under different sets of assumptions on e.g. the degree of price stickiness and then estimate single equations of the kind discussed in Section 2 on these simulated data. They find that depending on whether the shocks that affect the exchange rate are of a nominal or real nature, different variants of the basic PTM equation based on the discussion in Goldberg and Knetter (1997) perform differently. In particular, they find that versions of this 'standard' PTM equation that use better proxies for demand conditions perform better, in terms of bias, in estimating the ERPT coefficient when shocks are of a monetary nature. Versions using better measures of costs perform better in the presence of real shocks. In our study we have a richer specification compared to most of the literature, because we include in our sectoral VARs 'demand shifters', the price of domestic substitutes and measures of foreign costs, which we proxy by using not only foreign PPI but also global commodity prices. For

[14] We interpolate real GDP for the euro area with the index of industrial production applying the Chow and Lin (1971) procedure.

[15] We are aware of the problems related to using HP filtered variables, as exposed by Harvey and Jaeger (1993), but empirically this variable turned out systematically to have more explanatory power than the alternative measures of euro area internal demand that we experimented with.

this reason, the results of Corsetti, Dedola, and Leduc (2005) tentatively indicate that our approach is less likely to obtain distorted results than previous studies.

Marginal costs are the most relevant unobservable variable and foreign PPI is an admittedly rough proxy, since it reflects not only costs, but also mark-ups. On the other hand, our use of sector-specific PPI places us in a better position than most of the literature on sectoral ERPT, where only aggregate PPI variables, or even only CPI, are used.

The key variables for the sectors are shown in Figures 1 and 2 while those for the two aggregates are shown in Figure 3. Figure 4 displays the commodity price indices and the output gap series. Two main features of the data are rather clearly identifiable already by visual inspection in Figures 1 and 2: First, the import UVX series are much more volatile than the other series in all sectors. Second, while in some sectors, such as metals, the import UVX and the foreign and domestic prices have rather similar trends, in other sectors, like chemicals, such co-movements are not clearly visible. Figure 3 shows that the official import UVX series is smoother than the one we have constructed; this arises mostly from different computation methods, as Eurostat computes the official aggregate using the same source as we do, but creates unit values at the most disaggregated level, treats outliers at that level and uses a different aggregation procedure.[16] Also the nominal effective exchange rate series are quite volatile and do not exhibit a clear contemporaneous co-movement with the UVX series.

We next describe our modelling approach in more detail: $\Delta pcom$ is the rate of change of commodity prices in USD[17], gap is the euro area output gap, Δppi^* is the inflation rate of the effective foreign PPI, Δppi is the inflation rate of the euro area PPI, $\Delta neer$ is the rate of change of the effective nominal exchange rate and Δuvx is the rate of change of the unit value indices

[16] To be precise, the procedure, as described in Eurostat's manual on external trade statistics, is the following: monthly raw data are processed at the most detailed level in order to calculate elementary unit-values defined by trade value/quantity. These unit-values are divided by the average unit-value of the previous year to obtain elementary unit-value indices, from which outliers are detected and removed. Elementary unit-value indices are then aggregated over countries and commodities, by using the Laspeyres, Paasche and Fisher formulae. Finally, the Fisher unit-value indices are chained back to the reference year (2000=100) and are used to approximate the import and export price movements. Value-indices are calculated as the percentage change between the trade value of the current month and the average monthly trade value of the previous year.

[17] As indicated above, for some sectors we use specific commodity prices instead of or in addition to the oil price. For the discussion here we use for simplicity just some commodity price index and use the output gap as demand variable. Throughout the paper lower case letters for variables denote logarithms, except for the *gap*.

Figure 1: The nominal effective exchange rate (NEER, blue dotted line), foreign producer price index (PPI*, green dashed line), euro area producer price index (PPI, red continuous line) and unit value index (UVX, black continuous line) for the sectors investigated in this study.

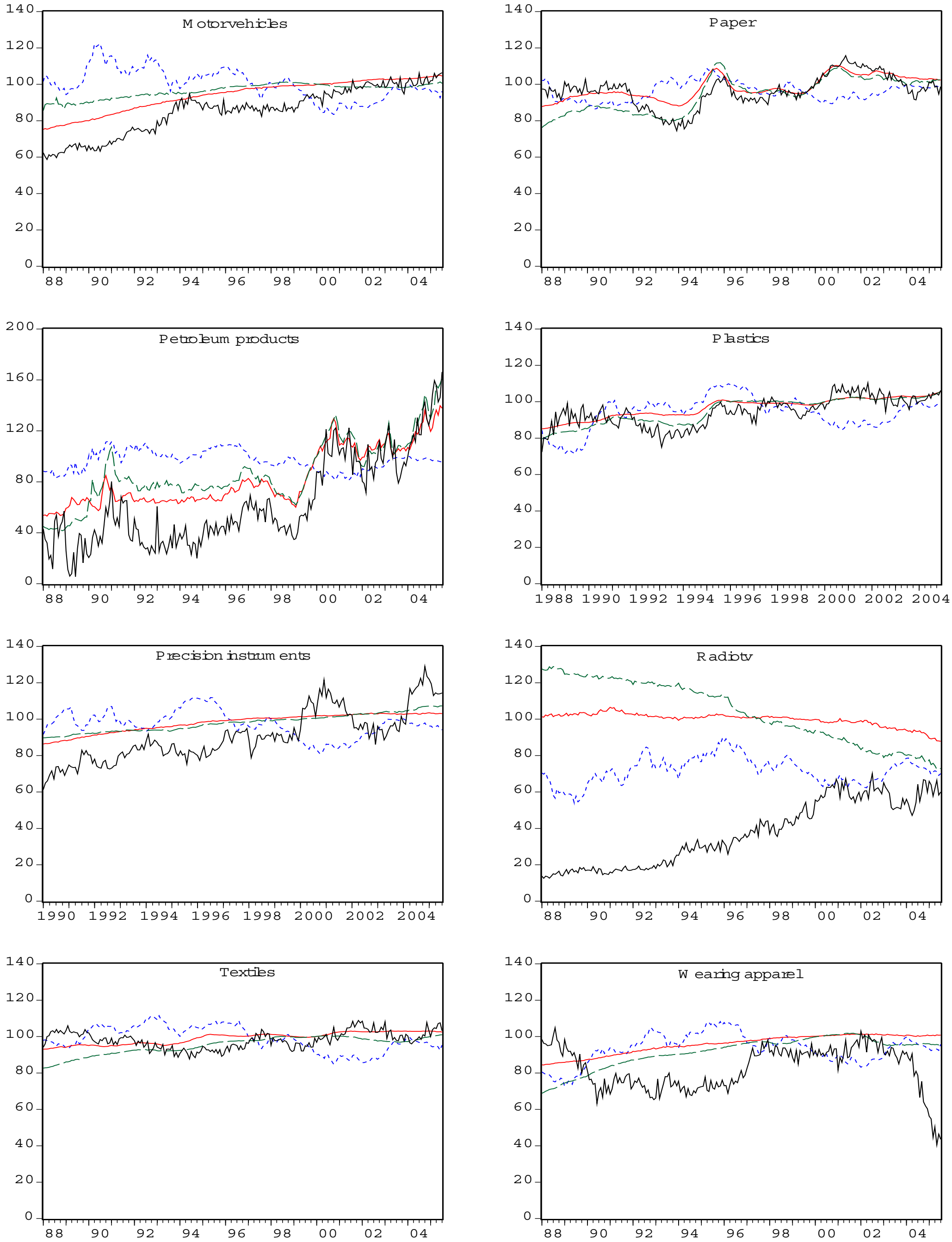

Figure 2: The nominal effective exchange rate (NEER, blue dotted line), foreign producer price index (PPI*, green dashed line), euro area producer price index (PPI, red continuous line) and unit value index (UVX, black continuous line) for the sectors investigated in this study.

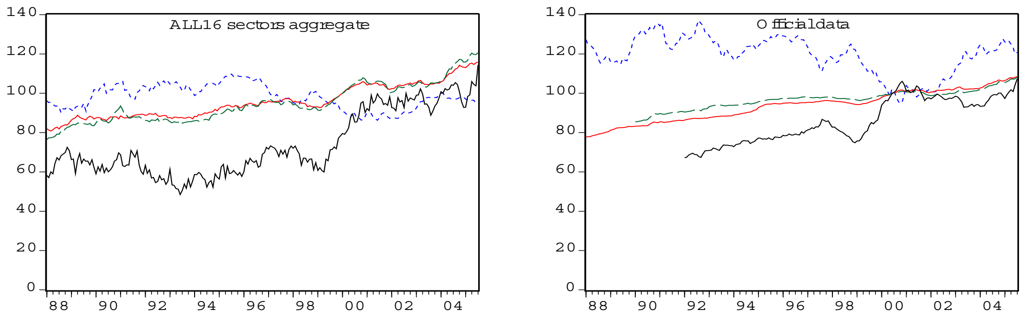

Figure 3: The nominal effective exchange rate (NEER, blue dotted line), foreign producer price index (PPI*, green dashed line), euro area producer price index (PPI, red continuous line) and unit value index (UVX, black continuous line) for the two manufacturing aggregates investigated in this study. The left figure shows the data obtained with our aggregation scheme outlined in the text and the right figure displays the 'official' data.

of imports. The joint vector is labelled as $\Delta y = \begin{bmatrix} \Delta pcom & gap & \Delta ppi^* & \Delta ppi & \Delta neer & \Delta uvx \end{bmatrix}'$.

Using the rates of change of the variables circumvents unit root nonstationarity issues (see Figures 1 to 4) but of course inhibits us from performing structural vector error correction model (VECM) analysis. We do this because cointegration analysis with the logarithms of the level variables has delivered only weak and mixed evidence for 'interpretable' cointegrating relationships and, in some sectors, no evidence of cointegration at all. Furthermore, we do not have a particular structural theoretical model underlying our empirical implementation that generates the type of restrictions required for structural VECM modelling in mind. As a consequence, we focus on VARs in first differences of logarithms (i.e. in growth rates) with identification only achieved by Cholesky decompositions of the reduced form errors.[18]

[18] Subsequent work will impose more structure on VARs in growth rates to disentangle the shocks in a more sophisticated way, such as Bernanke (1986), Sims (1992) or in particular as Kim and Roubini (2000). In the present paper, we are more concerned with generating stylized facts and with their robustness e.g. with respect to the Cholesky ordering. Consequently, the results of the present paper can also be interpreted as first-stage input into subsequent work. For other studies of ERPT based on VAR models see e.g. Faruqee (2004), Hahn (2003) and Ito, Sasaki, and Sato (2005) who all use Cholesky decompositions, McCarthy (2000) and Shambaugh (2005) who use Blanchard-Quah type restrictions or Hüfner and Schröder (2003) who use a VECM model. None of these studies, however, looks at sectorally disaggregated data.

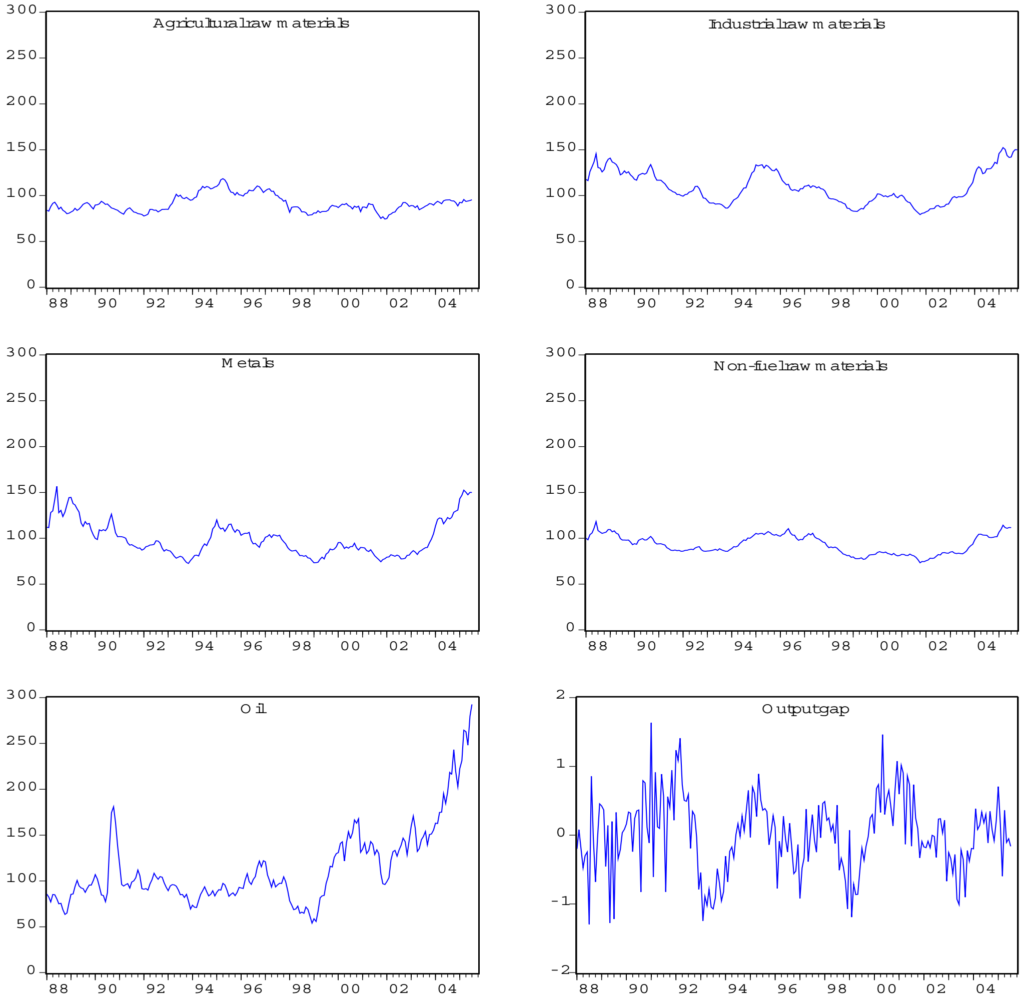

Figure 4: The additional variables used for modelling. All price indices are denominated in USD and are transformed to logarithmic differences in the estimations. The output gap is used without further transformations.

In this paper the structural form of a VAR for Δy is given by[19]

$$B\Delta y_t = b(L)\Delta y_t + u_t \tag{5}$$

where B is a regular matrix with (due to pure normalization) all diagonal entries equal to 1 that describes the contemporaneous links between the variables contained in Δy_t, $b(L) = \sum_{j=1}^{k} b_j L^j$ is a polynomial in the lag operator L and u_t, the vector of structural shocks, is a white noise process with diagonal covariance matrix Ω.[20]

Thus, in the structural form of a VAR each of the variables is explained by its own lags and by contemporaneous and lagged values of all other explanatory variables. Hence, e.g. the last equation of the equation system (5) reads as

$$\Delta uvx_t = \sum_{j=1}^{5} b_{6j}\Delta y_{t,j} + b_6(L)\Delta y_t + u_{t,6} \tag{6}$$

where b_{6j} denotes the corresponding entries of the sixth row of B, $b_6(L)$ denotes the sixth row of $b(L)$ and $u_{t,6}$ denotes the sixth component of u_t. It is immediate to see that equation (6) corresponds to the typical equation estimated in the single equation ERPT literature (up to lag length selection and specification issues).

As is well known, the structural form is not identified and hence cannot be estimated itself. Instead only the reduced form can be estimated from which *(over-)identified* structural forms can be recovered, see below. In our setting the reduced form is given by

$$\begin{aligned}\Delta y_t &= B^{-1}b(L)\Delta y_t + B^{-1}u_t &(7)\\ &= a(L)\Delta y_t + \varepsilon_t &(8)\end{aligned}$$

with $a(L) = B^{-1}b(L)$ and $\varepsilon_t = B^{-1}u_t$ with covariance matrix $\Sigma = B^{-1}\Omega(B^{-1})'$. If no coefficient restrictions are placed on (7) this equation system can be estimated equation by equation with OLS. In case of normally distributed errors this is efficient (in the sense of equivalent to maximum likelihood). The estimate for the covariance matrix is obtained from the variances and covariances computed from the individual equation residuals. OLS estimation works for the reduced form since no contemporaneous regressors are contained in

[19] For simplicity we neglect deterministic components like the intercept (always contained) throughout the discussion.

[20] To allow for a stationary solution of this linear dynamic equation the matrix polynomial $B - b(z)$ has to have its determinantal roots outside the closed unit circle.

any of the equations.[21] If restrictions are placed on the coefficients estimation of (7) can proceed by feasible GLS, which is a simple two-step procedure where the OLS residuals are used to compute an estimate of the variance covariance matrix of the residuals. This is relevant if one uses e.g. model reduction strategies as described in Lütkepohl and Krätzig (2004) or for hypothesis testing, which both generally lead to different regressors across equations.[22] Lag length selection is based on `BIC`.

For a system with n variables the matrix B contains $n \times (n-1)$ free parameters. Additionally there are n unknown variances in Ω. Thus, in total n^2 coefficients are unknown in (B, Ω). If we rely only upon contemporaneous identification there are however at most $\frac{n\times(n+1)}{2}$ coefficients that can be determined from Σ using the relationship

$$\Sigma = B^{-1}\Omega(B^{-1})'. \tag{9}$$

This implies that at least $\frac{n\times(n-1)}{2}$ restrictions have to be imposed on the pair (B, Ω). For a detailed exposition of this type of nonlinear identification problem arising in structural VARs, see Amisano and Giannini (1997).

The conceivably simplest way, originally proposed by Sims (1980), is to use the Cholesky decomposition of Σ to achieve exact identification. For every positive definite symmetric matrix Σ there exists a unique decomposition of the form $\Sigma = PP'$, with P lower triangular. This implies

$$PP' = B^{-1}\Omega^{1/2}\Omega^{1/2}(B^{-1})' \tag{10}$$

and hence $P = B^{-1}\Omega^{1/2}$ and consequently $B = \Omega^{1/2}P^{-1}$. If the variance of the structural shocks is assumed to equal the variance of the reduced form shocks this uniquely identifies both B and Ω.[23] Using the Cholesky decomposition imposes a contemporaneous causal ordering on the variables (and the shocks), since current values of variables only depend upon current values of variables that are 'above' that variable in the equation system.

There are two major drawbacks of this simple identification scheme: First, for a system of

[21] Of course, restricting B to be the identity matrix leads to an exactly identified structural form that coincides with the reduced form. From this extreme example we see that placing a sufficient number of restrictions on B can be used to solve the identification problem.

[22] Efficiency (with normally distributed errors) of GLS follows from the fact that a VAR with restrictions is a seemingly unrelated regression system.

[23] It is completely equivalent to assume that the structural form shocks have unit variance (i.e. $\Omega = I$), which implies that the matrix B is not normalized to have 1's on the diagonal. None of the interpretations changes when choosing one version or the other.

n variables there are $n!$ possible orderings, which e.g. means that for a 6 variable system there are 720 possible orderings.[24] Thus, 'preferred' orderings have to be chosen, if possible, by resorting to the economic theory underlying the problem at hand. Often, however, economic theory does not provide sufficiently clear and unambiguous guidance for this choice. In such a situation, Sims (1981) suggests validating the robustness of the results across different orderings. Second, from a statistical perspective all the possible orderings are equally valid descriptions of the data, since they result in the same value of the (pseudo) likelihood.[25]

Lacking a fully specified theoretical model that would lead to a unique (over-)identified structural form to assess the robustness of our findings we report results from placing the nominal effective exchange rate at different positions in the VAR.

Up to now we have treated all variables as endogenous, which we identified for our problem as a key advantage of the VAR framework. Also Sims (1980) sees this as a key advantage as opposed to imposing what he calls 'incredible' exogeneity assumptions. The VAR framework, however, equips us also to test for exogeneity of variables or groups of variables via simple zero-restriction coefficient tests. In fact, for all but two of our sectors and the aggregate variables the oil price and the output gap are found to be exogenous by (block) exogeneity testing. Thus, partition our vector Δy in two sub-blocks $\Delta y^1 = \begin{bmatrix} \Delta oil & gap \end{bmatrix}$ and $\Delta y^2 = \begin{bmatrix} \Delta ppi^* & \Delta ppi & \Delta neer & \Delta uvx \end{bmatrix}$ and consider $a(L)$ to be partitioned accordingly. Then (block-)exogeneity of Δy^1 can be tested by testing the null hypothesis that the coefficients in the upper right 2×4 corners (corresponding to the coefficients of Δy^2_{t-j} in the equations for Δy^1_t) of all coefficient matrices a_j for $j = 1, \ldots, k$ are jointly equal to zero. This can be tested by a simple likelihood ratio type test and estimation under the null hypothesis can be performed with feasible GLS. Sometimes this is also referred to as (multivariate) Granger non-causality. Since this hypothesis is not rejected for our data (with the exception of two sectors where the output gap appears to be endogenous) we treat those two variables as exogenous and have thus reduced the dimension of the VARs form six to four with two exogenous variables. Lag length selection for the exogenous variables is based on a testing sequence of non-significance of the highest lag until this hypothesis is rejected for the first time.

[24] Of course, estimation only has to be performed once, because the coefficients for any ordering can be computed by appropriate permutation operations on the rows of the coefficient matrices.

[25] This is also immediately clear from the observation that these are all exactly identified models, whereas hypothesis testing to discriminate amongst models requires *over-identifying* restrictions.

Completely analogously we subsequently also test for exogeneity of the remaining endogenous variables. Remember from the above discussion that single equation OLS estimation of the Δuvx equation is consistent only if the three variables Δppi^*, Δppi and $\Delta neer$ are all exogenous.

The dynamic effects of changes in the structural shocks are discussed within a VAR framework by investigating the impulse response functions, which are defined as $\frac{\partial \Delta y_{t+h,j}}{\partial u_{t,i}}$ for $h \geq 0$ and $i, j = 1, \ldots, n$. Under the assumption that the VAR process is stationary, the impulse response functions can be obtained from inverting the estimated autoregressive polynomial $a(L)$. Thus,

$$\begin{aligned}
\Delta y_t &= a^{-1}(L)\varepsilon_t \\
&= c(L)\varepsilon_t \\
&= c(L)B^{-1}u_t \\
&= \sum_{m=0}^{\infty} c_m P\Omega^{-1/2}u_{t-m} \qquad (11)
\end{aligned}$$

denoting with $c(L) = a^{-1}(L)$, which exists due to the stationarity assumption. Thus, the effect on variable $\Delta y_{t+h,j}$ emanating from a change of the structural form shock $u_{t,i}$ of size $\Delta u_{t,i}$ is given by the j-th entry of $c_h P\Omega^{-1/2}[0, \ldots, 0, \Delta u_{t,i}, 0, \ldots, 0]'$. The total effect on $\Delta y_{t+h,j}$ due to this shock is given by the sum of the effects from period 0 to h, i.e. the by the j-th entry in $\sum_{m=0}^{h} c_m P\Omega^{-1/2}[0, \ldots, 0, \Delta u_{t,i}, 0, \ldots, 0]'$, which is known as accumulated impulse response function. The so called long-run (accumulated) effect is given by the sum up to $h = \infty$.[26] In the empirical results below we use as our ERPT estimates at horizon h the accumulated impulse response function of the variable Δuvx to a unit structural shock in the equation for $\Delta neer$ up to this horizon.[27] Considering shocks of unit size allows for an interpretation in terms of elasticities or percentage changes in the import prices due to a one percent change in the nominal effective exchange rate. Due to the fact that the series coefficients c_m converge to zero at a geometric rate, the accumulated impulse response function will numerically flatten out at small, finite values of h already. Numerically it turns out that $h = 24$ is large enough,

[26] The long-run effect can also be computed more easily by noting that it is given by the corresponding entry in $c(1)P\Omega^{-1/2}[0, \ldots, 0, \Delta u_{t,i}, 0, \ldots, 0]'$ since $c(1) = \sum_{j=0}^{\infty} c_j$. Noting next that $c(1) = a^{-1}(1)$ facilitates computation further.

[27] Of course it is also interesting to study other effects, e.g. the impact of shocks to foreign PPI on UVX. We abstain from doing so in this Cholesky decomposition based paper and will investigate the effects of other shocks in subsequent work based on more structural models.

i.e. the dynamic response is essentially completed after two years (actually even faster).

Being based on estimated coefficient matrices, the numerical impulse response functions are themselves estimated quantities. Therefore, we display them together with their confidence bounds computed by bootstrapping as outlined in Hall (1992).

4 Estimation Results

The results of the VAR model specification analysis are reported in Tables 3 and 4.[28] As indicated above, model selection starts by lag length selection according to `BIC`, whose indications are complemented by residual-based specification analysis (like residual normality or no serial correlation in the residuals). The models presented in the table are well specified according to these measures.

Given the discussed interest in assessing stability of ERPT we investigate the structural stability of our estimated models by performing both the CUSUM and the CUSUMSQ tests, see Lütkepohl and Krätzig (2004). The CUSUM test does not indicate any structural instabilities at all and also the CUSUMSQ test delivers only very marginal evidence for instabilities arising in some of the sectors in the equations for Δppi^* (most notably in the computing and office machinery sector). The instabilities appear minor and additional efforts did not improve the modelling results. Therefore, we conclude that from a multivariate perspective structural instability of estimated ERPT is not an issue, in particular since the equation for Δuvx appears well specified for all sectors and both aggregates.[29] This finding is in contrast with some of the single-equation based evidence on ERPT instability, such as that reported in Otani, Shiratsuka, and Shirota (2003) and Marazzi et al. (2005). However, our sample starts in 1988, making our results not indicative of changes that might have taken place in the early to mid 1980's.

In Table 5 we report the results for the exogeneity tests for each of the four main endogenous variables in the specified VAR models. This table reports e.g. in the first column the results of testing the hypothesis that Δppi^* is exogenous for the remaining three variables Δppi, $\Delta neer$ and Δuvx.[30] The results in the table show that only in some of the sectors some

[28] Strictly speaking we specify so called VARX models, since we include exogenous variables.

[29] In two or three cases the test statistic marginally crosses the critical lines for a few time points.

[30] Note that we report here the test results testing each of the variables. For the single equation analysis of ERPT also the hypothesis that Δppi^*, Δppi and $\Delta neer$ are exogenous for Δuvx is relevant. This hypothesis is rejected except for four sectors discussed in the text.

Sector	Endogenous Variables	Lags	Exogenous variables	Lags	Sample
CHEM	Δppi^*, Δppi, $\Delta neer$, Δuvx	5	Δoil, gap	1	88:2 to 04:12
COMP	Δppi^*, Δppi, $\Delta neer$, Δuvx	5	Δoil, Δind, gap	2	90:2 to 04:12
ELEC	Δppi^*, Δppi, $\Delta neer$, Δuvx	5	Δoil, Δind, gap	1	90:2 to 04:12
FABMET	Δppi^*, Δppi, $\Delta neer$, Δuvx	3	Δmet, gap	1	88:2 to 04:12
FOODBEV	Δppi^*, Δppi, $\Delta neer$, Δuvx, gap	8	–	-	90:2 to 04:12
MACH	Δppi^*, Δppi, $\Delta neer$, Δuvx	6	Δoil, Δind, gap	2	88:2 to 04:12
METALS	Δppi^*, Δppi, $\Delta neer$, Δuvx	4	Δoil, Δind, Δmet, gap	2	88:2 to 04:12
MIN	Δppi^*, Δppi, $\Delta neer$, Δuvx	6	Δoil, gap	2	88:2 to 04:12
MOTOR	Δppi^*, Δppi, $\Delta neer$, Δuvx	6	Δoil, Δind, Δmet, gap	1	88:2 to 04:12

Table 3: Results of VAR model specification. All series used in logarithmic differences except for the output gap. Δmet is the rate of change of metals prices, Δind is the rate of change of industrial raw material prices and $\Delta nfuel$ is the rate of change of non-fuel commodity prices.

Sector	Endogenous Variables	Lags	Exogenous variables	Lags	Sample
PAPER	Δppi^*, Δppi, $\Delta neer$, Δuvx	4	gap	1	88:2 to 04:12
PETR	Δppi^*, Δppi, $\Delta neer$, Δuvx	8	-	-	92:2 to 04:12
PLAST	Δppi^*, Δppi, $\Delta neer$, Δuvx	6	Δoil, gap	2	88:2 to 04:12
PRECINSTR	Δppi^*, Δppi, $\Delta neer$, Δuvx	6	Δoil, Δmet, gap	1	90:2 to 04:12
RADIOTV	Δppi^*, Δppi, $\Delta neer$, Δuvx	4	Δoil, Δind, Δmet, gap	2	88:2 to 04:12
TEXTIL	Δppi^*, Δppi, $\Delta neer$, Δuvx, gap	6	Δagr	1	88:2 to 04:12
WEAR	Δppi^*, Δppi, $\Delta neer$, Δuvx	5	Δagr, gap	0	88:2 to 04:12
ALL16	Δppi^*, Δppi, $\Delta neer$, Δuvx	8	Δoil, gap	1	88:2 to 04:12
OFFICIAL	Δppi^*, Δppi, $\Delta neer$, Δuvx	5	Δoil, Δind, $\Delta nfuel$, gap	2	92:2 to 04:12

Table 4: Results of VAR model specification. All series used in logarithmic differences except for the output gap. Δmet is the rate of change of metals prices, Δind is the rate of change of industrial raw material prices and $\Delta nfuel$ is the rate of change of non-fuel commodity prices.

of these four variables are exogenous. In four sectors (COMP, MIN, MOTOR and RADIOTV) each of the three variables Δppi^*, Δppi and $\Delta neer$ is exogenous. For these four sectors also joint exogeneity of these three variables is not rejected.[31] Counting also the two aggregates, exogeneity necessary for consistent single equation OLS estimation is supported only in four out of 18 cases. This implies that for the data set at hand single-equation OLS estimation (as reported for a quantification of the differences in Section 4.1) is likely to produce inconsistent estimates.

Summing up, we have already established two important observations in contrast to the evidence collected in part of the available literature. First, we do not find evidence of structural instabilities of ERPT. Second, we find clear violations of the exogeneity assumptions necessary for consistency of the widely-used single-equation OLS-based ERPT estimation.

Let us now turn to the ERPT results. In Figures 5 to 7 we display the accumulated impulse response functions and 95% bootstrap confidence bounds for the Cholesky ordering as given in Δy^2. In this ordering we place the nominal effective exchange rate 'after' the foreign and domestic producer price indices to allow the exchange rate to adjust to price differentials within the period. Placing the unit value indices below the nominal effective exchange rate allows for contemporaneous effects of exchange rate changes on import prices.

For this specification the sectoral results in Figures 5 and 6 are quite in line with expectations. As already mentioned, we find quite some evidence of heterogeneity of ERPT across sectors. For eight sectors there is evidence of 'complete' ERPT (defined here in the sense that the confidence bounds around the impulse response function point estimate contain -1 but exclude 0 after several initial periods): computing and office machinery, electrical machinery, fabricated metals, plastics, precision instruments, radio and tv equipment, textiles and wearing apparel. For two sectors 0 is contained and -1 is not contained (in the long-run) in the confidence bounds around the ERPT point estimates, these are chemical products and motor vehicles (marginally). For the machinery and metals sector long-run ERPT is 'incomplete' in the sense that the confidence bounds include neither 0 nor -1. For some sectors (food products and beverages, mineral products, paper products (marginally) and petroleum products) the confidence bounds encompass both 0 and -1. Especially wide confidence bounds are found for mineral products and petroleum products. This may well reflect the large impact of commodity prices on these product categories' prices (even after controlling with commodity

[31] However, even for those sectors the VAR and single equation estimates of ERPT differ substantially.

	Δppi*	Δppi	Δneer	Δuvx	Exogenous variables
CHEM	0.00	0.01	0.02	0.02	Δoil, gap
COMP	0.78	0.21	0.48	0.01	Δoil, Δind, gap
ELEC	0.00	0.13	0.82	0.00	Δoil, Δind, gap
FABMET	0.00	0.74	0.38	0.00	Δmet, gap
FOODBEV*	0.05	0.00	0.00	0.00	–
MACH	0.01	0.00	0.16	0.00	Δoil, Δind, gap
METALS	0.00	0.01	0.27	0.00	Δoil, Δind, Δmet, gap
MIN	0.10	0.09	0.63	0.00	Δoil, gap
MOTOR	0.13	0.19	0.21	0.04	Δoil, Δind, Δmet, gap
PAPER	0.00	0.15	0.36	0.00	gap
PETR	0.00	0.58	0.16	0.00	–
PLAST	0.00	0.31	0.40	0.00	Δoil, gap
PRECINSTR	0.09	0.00	0.42	0.04	Δoil, Δmet, gap
RADIOTV	0.06	0.53	0.37	0.10	Δoil, Δind, Δmet, gap
TEXTIL*	0.07	0.00	0.03	0.00	Δagr
WEAR	0.09	0.00	0.42	0.04	Δagr, gap
ALL16	0.07	0.01	0.02	0.01	Δoil, gap
OFFICIAL	0.00	0.17	0.30	0.00	Δoil, Δind, $\Delta nfuel$, gap

Table 5: p-values of individual variable exogeneity tests for the four endogenous variables Δppi^*, Δppi, $\Delta neer$ and Δuvx in the final VARs.

The superscript * indicates that the output gap is also modelled as an endogenous variable.

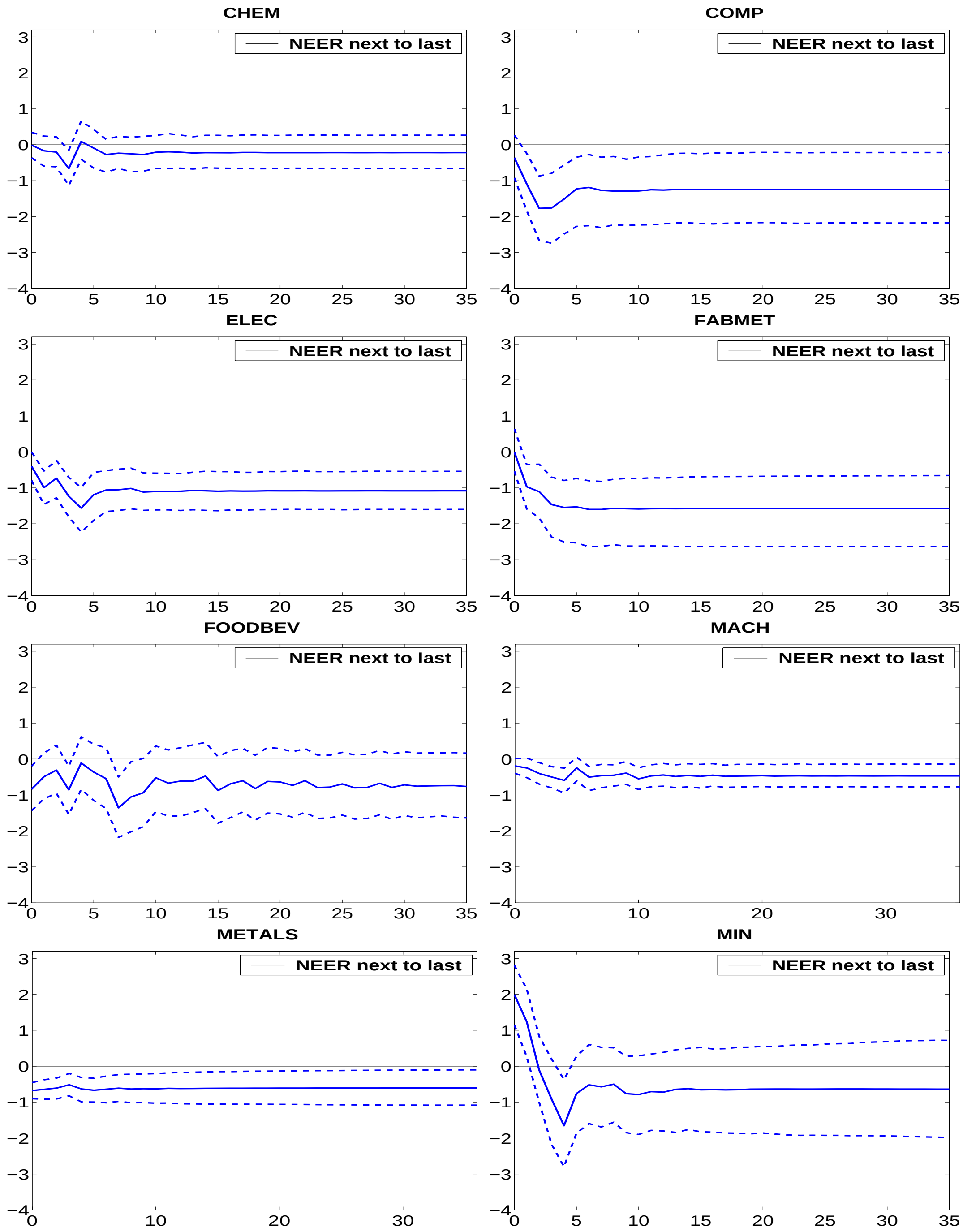

Figure 5: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$. The dashed lines are 95% bootstrap confidence bounds. The ordering of the endogenous variables is Δppi^{*}, Δppi, $\Delta neer$ and Δuvx.

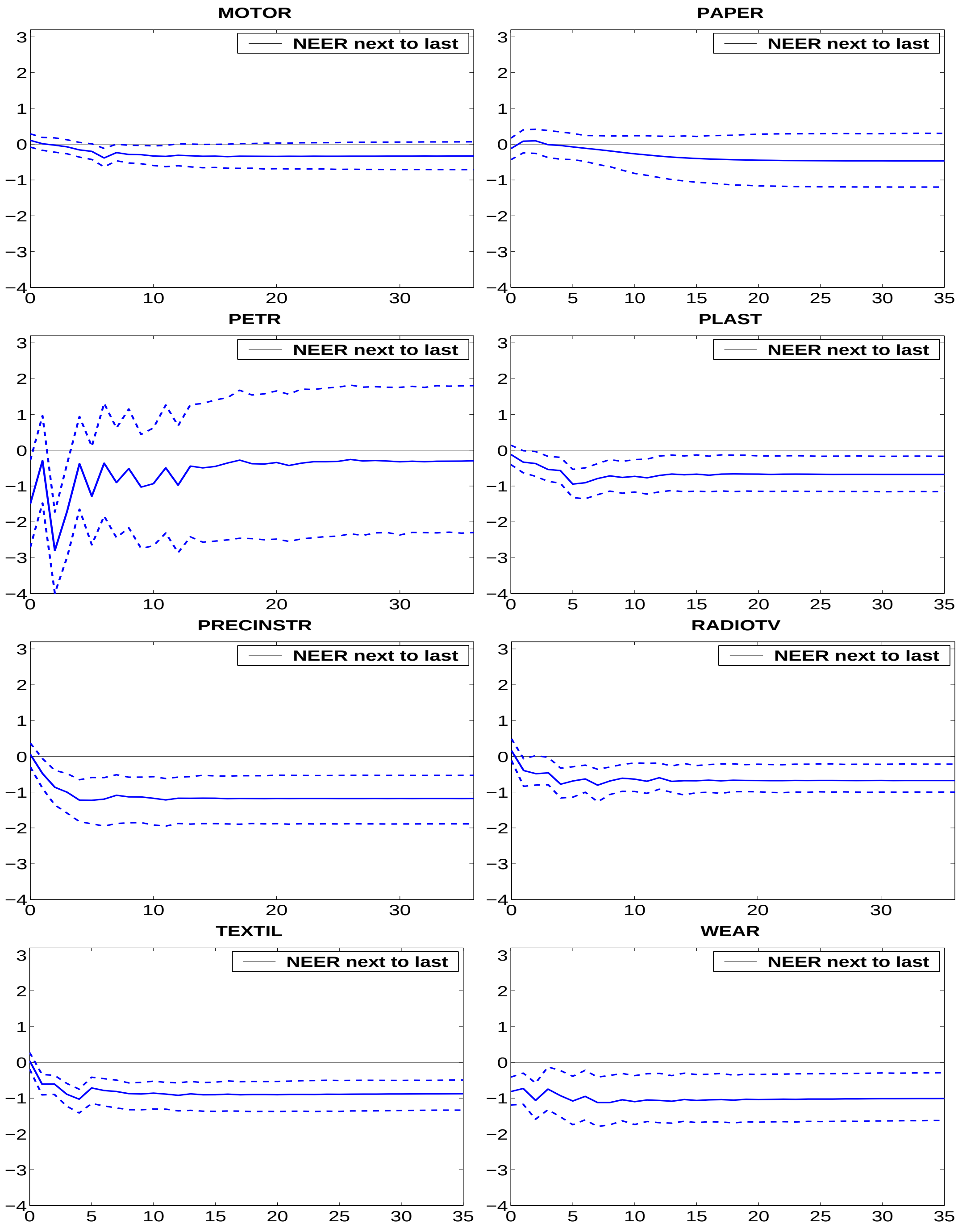

Figure 6: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$. The dashed lines are 95% bootstrap confidence bounds. The ordering of the endogenous variables is Δppi^*, Δppi, $\Delta neer$ and Δuvx.

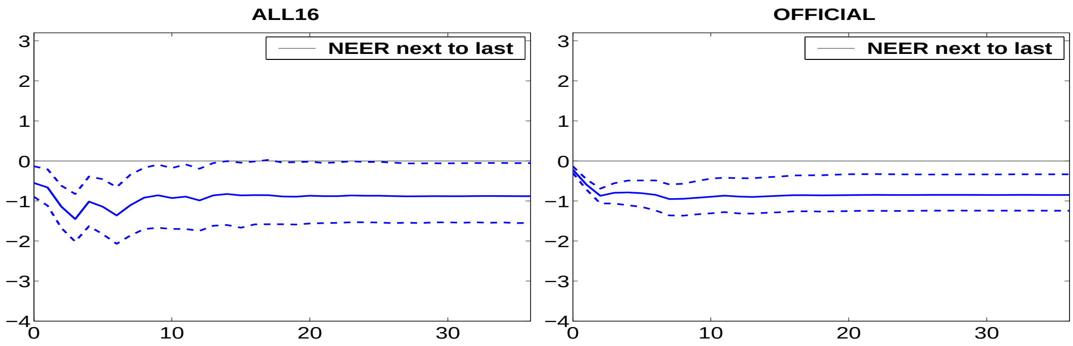

Figure 7: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$. The dashed lines are 95% bootstrap confidence bounds. The ordering of the endogenous variables is Δppi^*, Δppi, $\Delta neer$ and Δuvx.

price indices in the VAR models).

The shape of the impulse responses indicates rather fast adjustment, which is basically completed after 10 to 12 months. However, the size of both the short-run ERPT and the long-run ($h = 24$) ERPT differ substantially across sectors. This is an interesting observation since it indicates that heterogeneous ERPT dynamics across sectors have to be taken into account when assessing the impacts of exchange rate changes on the euro area, see also the third column of results in Tables 6 and 7. For $h = 0$ the point estimates range from -0.81 for wearing apparel to 2 for minerals. For $h = 24$ ERPT varies between -1.57 for fabricated metals and -0.12 (not significantly different from 0) for food products and beverages. These quite substantial differences in both the shape and magnitude of the accumulated impulse response functions indicate that a sectoral analysis is indeed important for understanding ERPT in the euro area.

In Figure 7 we display the results for the aggregate over the 16 sub-sectors of manufacturing constructed as described above (left plot) and the 'official' aggregate (right plot). The long-run ERPT is quite similar for both measures, with -0.86 for the ALL16 aggregate and -0.85 for the official aggregate. However, for small values of h differences emerge and for the ALL16 aggregate contemporaneous ERPT is given by -0.55 and for the official aggregate it is given by -0.22. The smaller short-run response of the official aggregate could be due to the smoothed construction of this series, which is also reflected in the smaller confidence bounds. For both aggregates long-run ERPT is complete in the sense defined above.

Several other studies have also investigated ERPT in the euro area, however, only at the aggregate level and mainly with respect to consumer prices.[32] Faruqee (2004), who studies ERPT along the pricing chain using data from 1990 to 2002 also finds ERPT to import prices close to complete in the long-run in a VAR containing different prices like import prices, export prices, consumer prices and wages. He finds rather low ERPT to consumer prices. Hüfner and Schröder (2003) also study ERPT to consumer prices for several European countries and they find, as Faruqee (2004), low pass-through to consumer prices, but faster and larger response of import prices. Hahn (2003) has a similar aim and also studies ERPT to different aggregate prices for the euro area using VAR models and Cholesky decompositions. She only finds an ERPT to import prices of about one half, which is below our as well as others' estimates, which are closer to one (respectively minus one). However, similarly to our results she finds, by estimating her model over sub-periods, no evidence for instabilities.

What about the robustness of the results with respect to the Cholesky ordering, suggested as a check in case of lack of theoretical guidance by Sims (1981)? Since we are only interested here in the effects of shocks to $\Delta neer$ we only display the results of four orderings (as opposed to all 24 possible orderings with four variables). We display the results when $\Delta neer$ is placed at any of the four positions where the three other variables are throughout ordered as Δppi^*, Δppi and Δuvx (and $\Delta neer$ potentially in between two of these). The results are shown in table format for horizons $h = 0$ and $h = 24$ in Tables 6 and 7 and graphically in Figures 11 to 13 in Appendix B.2.

Let us start with $h = 0$. By construction, when $\Delta neer$ is placed below Δuvx the contemporaneous ERPT is equal to 0 (left part of fourth column). For the other three orderings we see that it makes hardly any difference where $\Delta neer$ is placed, with two exceptions. For the chemicals sector contemporaneous ERPT is estimated to be 0.32 when $\Delta neer$ is placed first, and is 0.05 and -0.01 when placed second and third respectively. Note however that none of these numbers is statistically significantly different from 0. For the volatile minerals sector ERPT is 0.22 when $\Delta neer$ is placed first and about 2 when placed second or third. For the other sectors and the aggregates only very small differences occur.

The same picture emerges for $h = 24$, where again only for the two mentioned sectors sizeable differences emerge. As is also expected some small differences occur to the case when $\Delta neer$ is ordered last, but again only in some cases. For the aggregate measures this is more

[32] Thus, these studies are not directly comparable to our study where we focus on the inter-sectoral variation at the first step of the pricing chain.

pronounced for our measure than for the official series.

The graphical evidence collected in Figures 11 and 13 shows that this pattern also holds true for all values of h. Thus, generally the findings are quite robust with respect to the Cholesky orderings (at least with respect to the subset of possible orderings discussed). We interpret this, in the spirit of Sims (1981), as supportive of our results.

Sector	$\Delta neer = 1$				$\Delta neer = 2$				$\Delta neer = 3$				$\Delta neer = 4$			
	h=0		h=24		h=0		h=24		h=0		h=24		h=0		h=24	
CHEM	0.32		-0.08		0.05		-0.24		-0.01		-0.22		0.00		-0.21	
	-0.03	0.64	-0.48	0.36	-0.23	0.40	-0.70	0.22	-0.36	0.34	-0.66	0.27	0.00	0.00	-0.66	0.19
COMP	-0.35		-1.21		-0.33		-1.19		-0.36		-1.25		0.00		-1.03	
	-0.94	0.28	-2.19	0.00	-0.92	0.31	-2.23	-0.10	-0.92	0.26	-2.19	-0.22	0.00	0.00	-1.93	-0.08
ELEC	-0.04		-0.14		-0.05		-0.14		-0.05		-0.14		0.00		-0.11	
	-0.10	0.01	-0.21	-0.06	-0.10	0.01	-0.21	-0.06	-0.10	0.00	-0.20	-0.07	0.00	0.00	-0.17	-0.05
FABMET	0.01		-1.68		-0.06		-1.78		0.00		-1.57		0.00		-1.57	
	-0.49	0.57	-2.73	-0.75	-0.53	0.46	-2.81	-0.85	-0.54	0.64	-2.63	-0.67	0.00	0.00	-2.41	-0.68
FOODBEV*	-0.12		-0.13		-0.12		-0.13		-0.12		-0.12		0.00		-0.05	
	-0.21	-0.03	-0.26	0.02	-0.20	-0.03	-0.26	0.00	-0.21	-0.03	-0.24	0.02	0.00	0.00	-0.18	0.08
MACH	-0.18		-0.46		-0.17		-0.45		-0.19		-0.46		0.00		-0.40	
	-0.39	0.06	-0.76	-0.10	-0.40	0.05	-0.79	-0.10	-0.40	0.01	-0.77	-0.13	0.00	0.00	-0.66	-0.09
METALS	-0.62		-0.93		-0.62		-0.88		-0.68		-0.61		0.00		-0.25	
	-0.84	-0.41	-1.39	-0.43	-0.82	-0.42	-1.30	-0.44	-0.90	-0.46	-1.07	-0.12	0.00	0.00	-0.74	0.22
MIN	0.22		-0.08		1.97		-0.61		2.00		-0.63		0.00		-1.81	
	0.13	0.32	-0.23	0.08	1.06	2.87	-2.04	0.79	1.15	2.81	-1.92	0.59	0.00	0.00	-3.49	-0.37
MOTOR	0.09		-0.39		0.09		-0.38		0.11		-0.34		0.00		-0.40	
	-0.12	0.27	-0.74	0.01	-0.10	0.28	-0.74	0.01	-0.08	0.29	-0.69	0.04	0.00	0.00	-0.74	-0.05

Table 6: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$ at horizons $h = 0$ and $h = 24$. Point estimates are displayed in the first rows and bootstrap 95% confidence bounds are given in the second rows.

$\Delta neer = i$, $i = 1, \ldots, 4$ indicates the position of $\Delta neer$ in the Cholesky chain, where the other three variables are throughout ordered as Δppi^*, Δppi and Δuvx. In rows with * position 3 and 4 refer respectively to second-to last and last in a five-variable VAR with the output gap included as additional endogenous variable ordered first.

Sector	$\Delta neer = 1$				$\Delta neer = 2$				$\Delta neer = 3$				$\Delta neer = 4$			
	h=0		h=24		h=0		h=24		h=0		h=24		h=0		h=24	
PAPER	-0.15		-0.55		-0.12		-0.40		-0.12		-0.46		0.00		-0.39	
	-0.48	0.11	-1.32	0.19	-0.44	0.19	-1.22	0.36	-0.43	0.17	-1.18	0.29	0.00	0.00	-0.99	0.33
PETR	-1.86		-0.95		-1.50		-0.26		-1.50		-0.32		0.00		0.21	
	-3.18	-0.48	-3.25	1.34	-2.70	-0.26	-2.66	1.95	-2.71	-0.28	-2.44	1.70	0.00	0.00	-2.32	2.56
PLAST	-0.06		-0.36		-0.11		-0.66		-0.12		-0.67		0.00		-0.41	
	-0.30	0.19	-0.96	0.19	-0.35	0.15	-1.17	-0.15	-0.40	0.14	-1.14	-0.15	0.00	0.00	-0.74	-0.09
PRECINSTR	0.07		-1.14		0.07		-1.15		0.06		-1.18		0.00		-1.22	
	-0.27	0.43	-1.85	-0.40	0.41	-0.31	-0.45	-1.91	-0.29	0.37	-1.89	-0.54	0.00	0.00	-1.88	-0.44
RADIOTV	0.17		-0.67		0.16		-0.67		0.16		-0.67		0.00		-0.74	
	-0.19	0.51	-1.11	-0.28	-0.10	0.49	-1.08	-0.18	-0.11	0.49	-0.99	-0.22	0.00	0.00	-1.16	-0.33
TEXTIL*	-0.05		-0.86		-0.05		-0.85		0.04		-0.89		0.00		-0.91	
	-0.29	0.22	-1.31	-0.48	-0.31	0.16	-1.31	-0.43	-0.19	0.27	-1.37	-0.51	0.00	0.00	-1.40	-0.49
WEAR	-0.87		-1.03		-0.81		-1.02		-0.81		-1.03		0.00		-0.52	
	-1.27	-0.48	-1.66	-0.40	-1.20	-0.43	-1.68	-0.40	-1.19	-0.41	-1.66	-0.32	0.00	0.00	-1.16	0.14
ALL16	-0.57		-0.87		-0.57		-0.85		-0.55		-0.86		0.00		-1.22	
	-0.98	-0.16	-1.56	0.11	-0.89	-0.24	-1.53	-0.01	-0.90	-0.13	-1.55	-0.01	0.00	0.00	-2.18	-0.24
OFFICIAL	-0.26		-0.95		-0.18		-0.65		-0.22		-0.85		0.00		-0.57	
	-0.34	-0.17	-1.37	-0.54	-0.23	-0.11	-0.93	-0.35	-0.30	-0.13	-1.25	-0.33	0.00	0.00	-0.98	-0.18

Table 7: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$ at horizons $h = 0$ and $h = 24$. Point estimates are displayed in the first rows and bootstrap 95% confidence bounds are given in the second rows.
$\Delta neer = i$, $i = 1, \ldots, 4$ indicates the position of $\Delta neer$ in the Cholesky chain, where the other three variables are throughout ordered as Δppi^*, Δppi and Δuvx. In rows with * position 3 and 4 refer respectively to second-to last and last in a five-variable VAR with the output gap included as additional endogenous variable ordered first.

As a final robustness experiment we compare in Figure 8 the findings for the motor vehicles sector for two specifications. In the left graph we display again the results from the four endogenous (and three exogenous) variable VAR with $\Delta neer$ ordered third. In the right graph we display the results from a seven variable VAR (with the ordering of the variables given in the caption to the figure). In this seven endogenous variables model the three variables Δoil, Δind and gap are included in this order as further endogenous variables. To make the differences as large as possible we place $\Delta neer$ fourth in the seven variable VAR (but similar results are obtained even when placing $\Delta neer$ first). It turns out that the differences are minimal, especially at longer horizons. This indicates that the inefficiencies due to treating exogenous variables endogenous do not have a sizeable impact on the estimated dynamics and extent of ERPT, since similar findings also arise in other sectors.

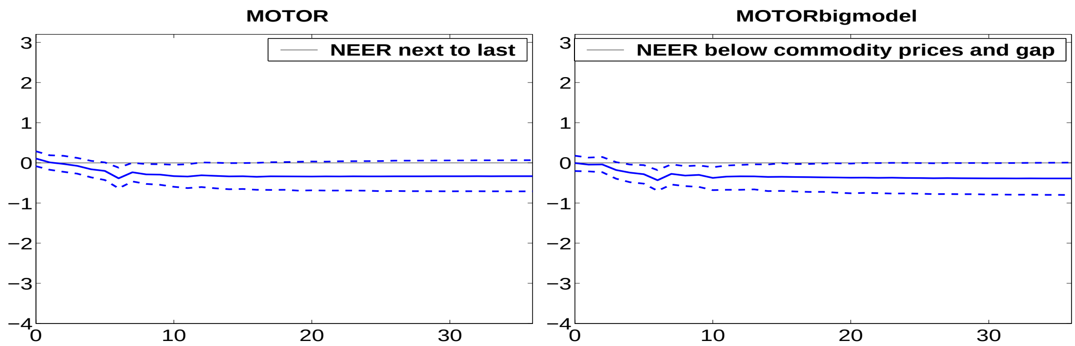

Figure 8: Comparison of accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$ across two different models. The left model corresponds to the one described above and the results are the same as displayed in the corresponding graph in Figure 5. The right model contains seven endogenous variables ordered as Δoil, Δind, gap, $\Delta neer$, Δppi^*, Δppi and Δuvx.

4.1 Single-equation results

In this section we compare our VAR results with results obtained by estimating equations of the form (6) by single equation OLS. Given that our VAR models appear well specified and nest our single equations (up to differences in the model selection procedure and the identification scheme) the differences in the estimated ERPT can be interpreted as the 'endogeneity bias' arising in the single-equation OLS framework.[33]

We present single-equation estimates based on a dynamic specification selected by PcGets in order to compare our well-specified VAR results with 'well-specified' single equation results. This appears to us as the relevant comparison exercise.[34] The details of the specification search algorithm implemented in PcGets are described in detail in Hendry and Krolzig (2001).[35] We start the model selection algorithm by taking as the general model an equation that allows for lagged values up to twelve of the dependent variable and contemporaneous and up to twelve lags of the explanatory variables. For the sake of comparability, we impose that the nominal exchange rate must be included at least at its contemporaneous value. The detailed results concerning the equations selected by PcGets are available from the authors upon request.

In Tables 8 and 9 we compare the results of the two methods. For comparison we choose the VAR results based on $\Delta neer$ ordered third (which are the results displayed also in Figures 5 to 7).

Let the estimated single equation be given by:

$$\Delta uvx_t = \sum_{j=1}^{J} \beta_{1,j} \Delta uvx_{t-j} + \sum_{k=0}^{K} \beta_{2,k} \Delta neer_{t-k} + \textbf{other regressors} + u_t \tag{12}$$

The (S-R) ERPT from the single equation displayed in the table is given by the coefficient estimate to the contemporaneous regressor $\Delta neer$, i.e. S-R $= \beta_{2,0}$. The long-run (L-R) ERPT estimate is given by

$$\text{L-R} = \frac{\beta_{2,0} + \cdots + \beta_{2,K}}{1 - (\beta_{1,1} + \cdots + \beta_{1,J})}.$$

In the table we also display the two standard error confidence bands for the single-equation estimates and again the 95% bootstrap confidence bounds for the VAR results. The results show substantial differences between the PcGets and VAR results for most sectors and smaller differences only for a few sectors. Differences occur in both the short-run and the long-run estimated ERPT but are generally larger for the latter, where for five sectors a different conclusion concerning the question whether ERPT is significantly different from 0 is reached

[33]To be more precise: Conditional upon the VAR models being correctly specified, this 'bias' is a finite-sample estimate of the endogeneity bias arising from the endogeneity of regressors in the single equation (12) (which impacts also on model selection) in the text.

[34]For completeness also single equation estimates with a similar specification as in the VARs are available upon request.

[35]For a critical theoretical analysis on the potential of model selection see Leeb and Pötscher (2005).

Sector	PcGets				VAR			
	S-R		L-R		h=0		h=24	
CHEM	-0.14		**-0.23**		-0.01		-0.22	
	-0.57	0.29	-0.43	-0.03	-0.36	0.34	-0.66	0.27
COMP	-0.23		**-0.96**		-0.36		**-1.25**	
	-0.92	0.46	-1.73	-0.19	-0.92	0.26	-2.19	-0.22
ELEC	**-0.60**		**-0.94**		**-0.40**		**-1.08**	
	-1.08	-0.11	-1.40	-0.49	-0.80	0.00	-1.60	-0.54
FABMET	0.20		-0.96		0.00		**-1.57**	
	-0.44	0.85	-1.44	-0.48	-0.54	0.64	-2.63	-0.67
FOODBEV	**-0.77**		-0.41		**-0.84**		-0.79	
	-1.39	-0.14	-0.96	0.13	-1.43	-0.20	-1.65	0.11
MACH	-0.02		**-0.39**		-0.19		**-0.46**	
	-0.30	0.27	-0.54	-0.24	-0.40	0.01	-0.77	-0.13
METALS	**-0.62**		**-1.56**		**-0.68**		**-0.61**	
	-0.83	-0.41	-1.91	-1.20	-0.90	-0.46	-1.07	-0.12
MIN	1.99		-0.15		1.77		-0.71	
	1.12	2.86	-0.99	0.69	0.98	2.49	-1.91	0.39
MOTOR	0.15		**-0.17**		0.11		-0.34	
	-0.04	0.34	-0.32	-0.02	-0.08	0.29	-0.69	0.04

Table 8: Comparison of single equation and VAR results with $\Delta neer$ at third position.
In the first rows *S-R* displays the contemporaneous effect and *L-R* displays the long-run effect (as described in the text) for the PcGets results and the VAR results are displayed at horizons $h = 0$ and $h = 24$.
In the second rows we display the two standard error confidence bounds for the single equation estimates and the 95% bootstrap confidence bounds for the VAR results. Estimates significantly different from zero are in bold.

Sector	PcGets				VAR			
	S-R		L-R		h=0		h=24	
PAPER	-0.15		-0.29		-0.12		-0.46	
	-0.44	0.14	-0.55	-0.03	-0.43	0.17	-1.18	0.29
PETR	**-1.44**		**-2.38**		**-1.50**		-0.32	
	-2.78	-0.09	-3.50	-1.25	-2.71	-0.28	-2.44	1.70
PLAST	-0.25		**-0.24**		-0.12		**-0.67**	
	-0.54	0.04	-0.39	-0.08	-0.40	0.14	-1.14	-0.15
PRECINSTR	0.02		**-0.49**		0.06		**-1.18**	
	-0.40	0.45	-0.92	-0.06	-0.29	0.37	-1.89	-0.54
RADIOTV	0.30		-0.11		0.16		**-0.67**	
	-0.09	0.69	-0.43	0.21	-0.11	0.49	-0.99	-0.22
TEXTIL	0.05		**-0.60**		0.04		**-0.89**	
	-0.22	0.33	-0.83	-0.37	-0.19	0.27	-1.37	-0.51
WEAR	-0.97		**-1.40**		**-0.81**		**-1.03**	
	-1.38	-0.56	-1.98	-0.83	-1.19	-0.41	-1.66	-0.32
ALL16	**-0.65**		**-0.49**		**-0.55**		**-0.86**	
	-1.23	-0.07	-0.94	-0.05	-0.89	-0.13	-1.52	-0.02
OFFICIAL	**-0.18**		**-0.60**		**-0.22**		**-0.85**	
	-0.28	-0.09	-0.73	-0.47	-0.30	-0.13	-1.25	-0.33

Table 9: Comparison of single equation and VAR results with $\Delta neer$ at third position.
In the first rows *S-R* displays the contemporaneous effect and *L-R* displays the long-run effect (as described in the text) for the PcGets results and the VAR results are displayed at horizons $h = 0$ and $h = 24$.
In the second rows we display the two standard error confidence bounds for the single equation estimates and the 95% bootstrap confidence bounds for the VAR results. Estimates significantly different from zero are in bold.

by applying the two approaches. Thus, indeed the discussed potential endogeneity bias is seen to substantially influence the findings with our data set. It is very likely that similar problems plague other single equation OLS studies.

5 Summary and Conclusions

In this paper we estimate ERPT to euro area manufacturing import prices at the sectorally disaggregated level as well as for the manufacturing aggregate. Using a VAR framework with additional (exogenous) explanatory variables allows to derive well specified dynamic models that take into account that some major explanatory variables (in particular domestic and foreign PPI, and the nominal effective exchange rate) are endogenous. Endogeneity of these variables can be assessed within the VAR framework by hypothesis testing and in general the aforementioned variables are found to be endogenous. This renders widely-used single-equation OLS estimates of ERPT inconsistent (due to the 'endogeneity bias'). As demonstrated in Section 4.1 the ERPT estimates based on our system estimates and based on single-equation OLS estimation differ substantially. We speculate that similar findings also hold true for the data sets used in single-equation OLS studies in the literature.

A second advantage of a VAR system approach is that it at allows to not only study the extent of ERPT, but also its dynamics. The dynamic measure of ERPT over h periods is given by the impulse response function of import prices to a shock in the nominal effective exchange rate accumulated over h periods. Using this quantity as our dynamic measure we find that in general ERPT adjustments are essentially completed after one year.

We find very heterogeneous results across sectors, for immediate responses, to a certain extent for the dynamics and also for the long-run response. Long-run ERPT (at $h = 24$) ranges from -0.12 (not significantly different from 0) for food and beverages to -1.57 (not significantly different from -1) for fabricated metals. These results point to the importance of studying ERPT at a disaggregated level to understand the inflationary impact of exchange rate changes. Similarly to a sectoral disaggregation, an intra-euro area country disaggregation may lead to sharpened insights concerning euro area price adjustment in response to exchange rate fluctuations.

For the two aggregate measures long-run ERPT is essentially identical, -0.85 and -0.86, and not significantly different from -1. However, the short-run ERPTs differ and short-run

ERPT is larger for our exactly aggregated variables. This may well reflect the fact our variables provide, due to the exact construction, a cleaner basis for assessing ERPT.

By analyzing the stability of our estimated models we contribute to the recent discussion concerning structural changes, in particular declines, of ERPT (compare the discussion in Sections 1 and 2). By means of CUSUM and CUSUMSQ tests we do not find evidence for structural instabilities and hence declining ERPT on the sectoral or the manufacturing aggregate levels.

Future work based on this study will be along three directions. The first is, as already indicated above to study ERPT disaggregated also across countries. Second, it will be important to study ERPT at a disaggregated level also to other prices, most notably export prices. Third, as mentioned in the text, more structural identification schemes will be used to shed light on the dynamics of import (respectively export) prices to several well-identified structural shocks.

Appendix A: Details of Data Set

Appendix A.1.: Sources and pre-treatment of data

The PPI series for the euro area countries are from Eurostat's STS database, while those for the partner countries are collected from different sources: data for European Union countries are available from Eurostat, while Global Insight data are used for the USA, Japan and Korea. US data for some sectors have to be back-cast using data from the BLS, because they are only available starting in 2003. For most countries, the data from OECD's Indicators of Industry and Services (IIS, available 1990-2001 in ISIC 3.1) and Indicators of Industrial Activity (IIA, available 1975-1998 in the older ISIC classification) are precious sources of back-data which are used to back-cast the shorter series via chain-linking. Unfortunately, data from the IIS were discontinued in 2001.

Data on exchange rates are readily available from IFS, BIS and, after 1999, from the ECB.

The import unit value indices are constructed based on import values in ECU-EUR and on import volumes in 1000Kg. The source is Eurostat's COMEXT database. The values are converted to national currencies before calculating the unit values, which are then indexed to a base period. Given the volatility of the import value and volume data, a detailed analysis of and correction for outliers is performed. Table 10 summarizes the data corrections. For all volume and value series, every observation x_t for which the standardized month-on-month growth rate: $\left(\frac{x_t}{x_{t-1}} - 1\right)$ is larger than 3 is set to the average of x_{t-12} and x_{t+12}. The final aggregates are then run through TRAMO-SEATS to identify remaining outliers and seasonal patterns, which are removed when present. In the table, the suffix_lin indicates that the series is adjusted for outliers by removing the irregular component identified by TRAMO. The suffix _sa indicates that the series is seasonally adjusted. An asterisk indicates that TRAMO is set to remove not only additive and temporary outliers, but also level shifts.

Table 11 reports the description and definition of the price variables used in this study.

	UVX	NEER	PPIstar	PPI	UVXea	Start
CHEM	uvx_lin_sa	neer	ppistar	ppi	uvxea_lin_sa	88:1
ELEC	uvx_missing_lin	neer_sa	ppistar_lin	ppi_lin_sa	uvxea_lin_sa	90:1
MOTOR	uvx_lin_sa	neer_lin	ppistar_sa	ppi_lin_sa	uvxea_lin_sa	88:1
MIN	uvx_lin_sa*	neer	ppistar_lin_sa*	ppi_lin_sa	uvxea_lin_sa	88:1
PETR	uvx	neer	ppistar	ppi	uvxea_lin	88:1
FOODBEV	uvx_lin_sa	neer	ppistar_lin	ppi_lin	uvxea_lin_sa	90:1
PAPER	uvx_lin	neer_lin_sa	ppistar_lin	ppi_lin_sa	uvxea_lin_sa	88:1
METALS	uvx_sa	neer	ppistar_lin	ppi_lin	uvxea_lin_sa	88:1
COMP	uvx_lin_sa*	neer	ppistar_lin*	ppi_lin_sa*	uvxea_lin_sa	90:1
TEXTIL	uvx	neer_sa	ppistar_lin_sa	ppi_lin_sa	uvxea_lin_sa	88:1
WEAR	uvx_lin_sa	neer_sa	ppistar_lin_sa	ppi_lin_sa	uvxea_lin_sa	88:1
FABMET	uvx_lin_sa	neer_sa	ppistar_lin_sa	ppi_lin_sa	uvxea_lin_sa	88:1
PRECINSTR	uvx_lin_sa	neer	ppistar_lin	ppi_lin_sa	uvxea_lin_sa	90:1
RADIOTV	uvx_lin_sa*	neer_lin*	ppistar_lin_sa*	ppi_lin_sa	uvxea_lin_sa	88:1
MACH	uvx_lin	neer	ppistar_lin	ppi_lin	uvxea_lin_sa	88:1
PLAST	uvx_lin	neer_lin	ppistar_lin	ppi_lin_sa	uvxea_lin_sa	88:1
ALL16	uvx_lin_sa	neer	ppistar_lin	ppi_lin	uvxea_sa	88:1
OFFICIAL	uvx	neer	ppistar_lin_sa	ppi_lin_sa	–	92:1

*Level shift also allowed

Table 10: Summary of data adjustments prior to econometric analysis.

Abbreviation	**Description**	Source and code (when available)
ALL	All commodities price	IMF, IFS.M.00176ACDZF...
Energy	Energy commodities price	IFS.M.00176ENDZF...
nonFuel	Non-fuel commodities price	IFS.M.00176NFDZF...
Oil	Average world price of crude oil	IFS.M.00176AAZZF...
Metals	Metals price	IFS.M.00176AYDZF...
AgrRawMaterials	Agricultural raw materials price	IFS.M.00176BXDZF...
Beverages	Beverages price	IFS.M.00176DWDZF...
Food	Food price	IFS.M.00176EXDZF...
Ind raw mat	Industrial raw materials price	-
GDP	EA GDP	OECD, MEI
IIP	EA index of industrial production	STS
M3	EA broad monetary aggregate	ECB and constructed based on EA averages
Int3	3-month nominal interest rate	ECB and constructed based on EA averages

Table 11: Definitions and sources of additional explanatory variables used.

Appendix A.2.: Weighting scheme and description of weights

(to be included or made available upon request)

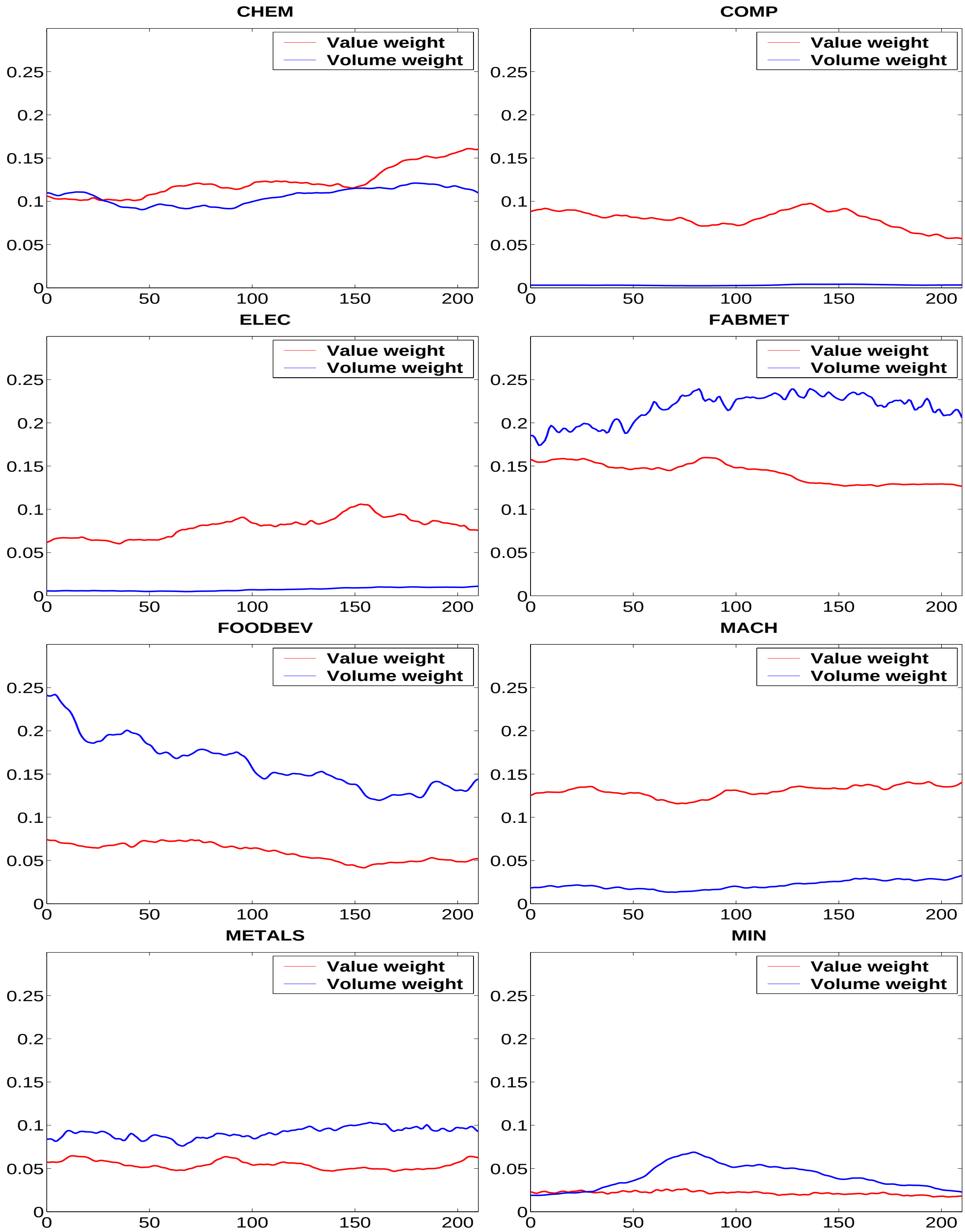

Figure 9: Sector weights, based on values and on volumes, smoothed via TRAMO-SEATS.

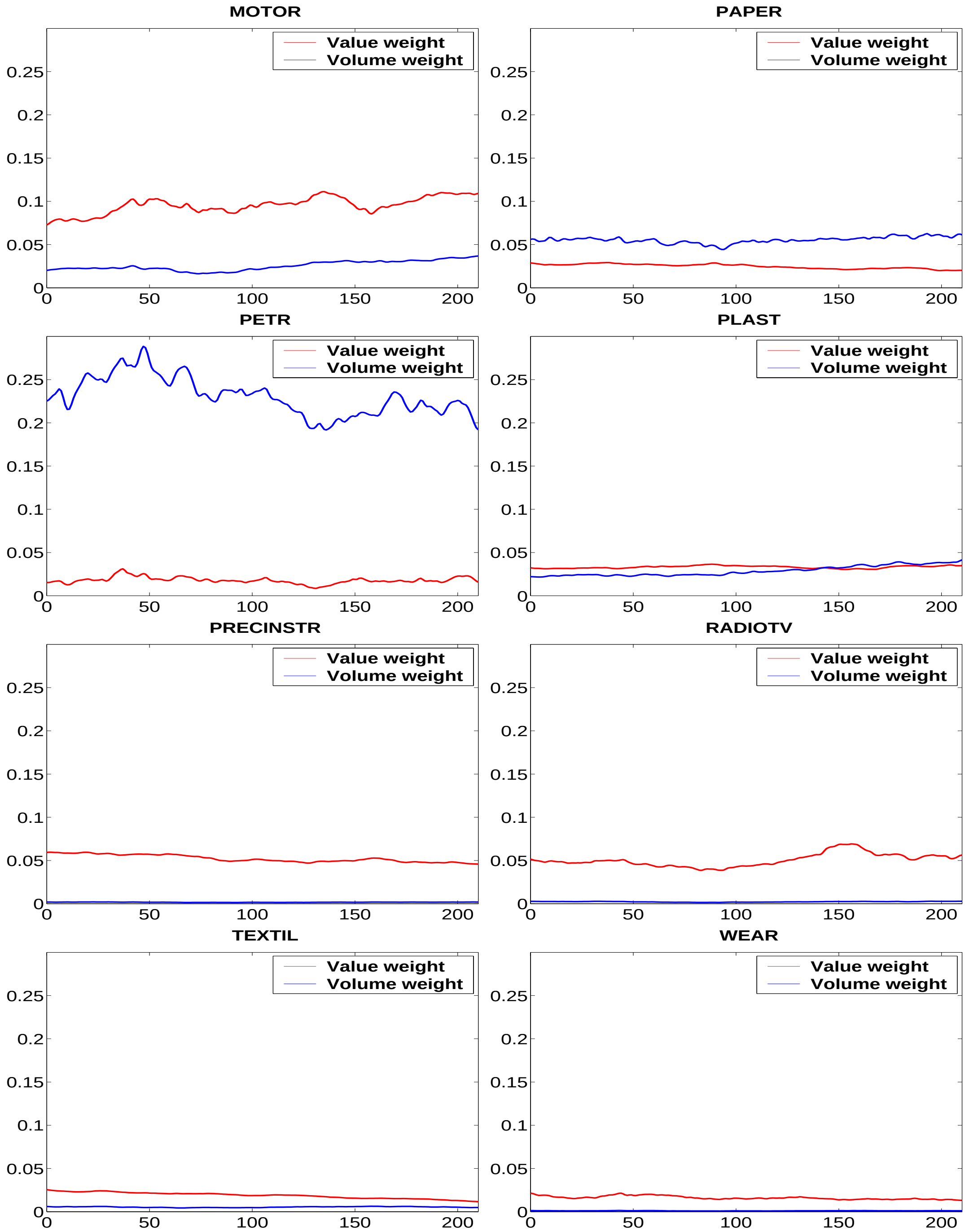

Figure 10: Sector weights, based on values and on volumes, smoothed via TRAMO-SEATS.

Appendix B: Additional Empirical Results

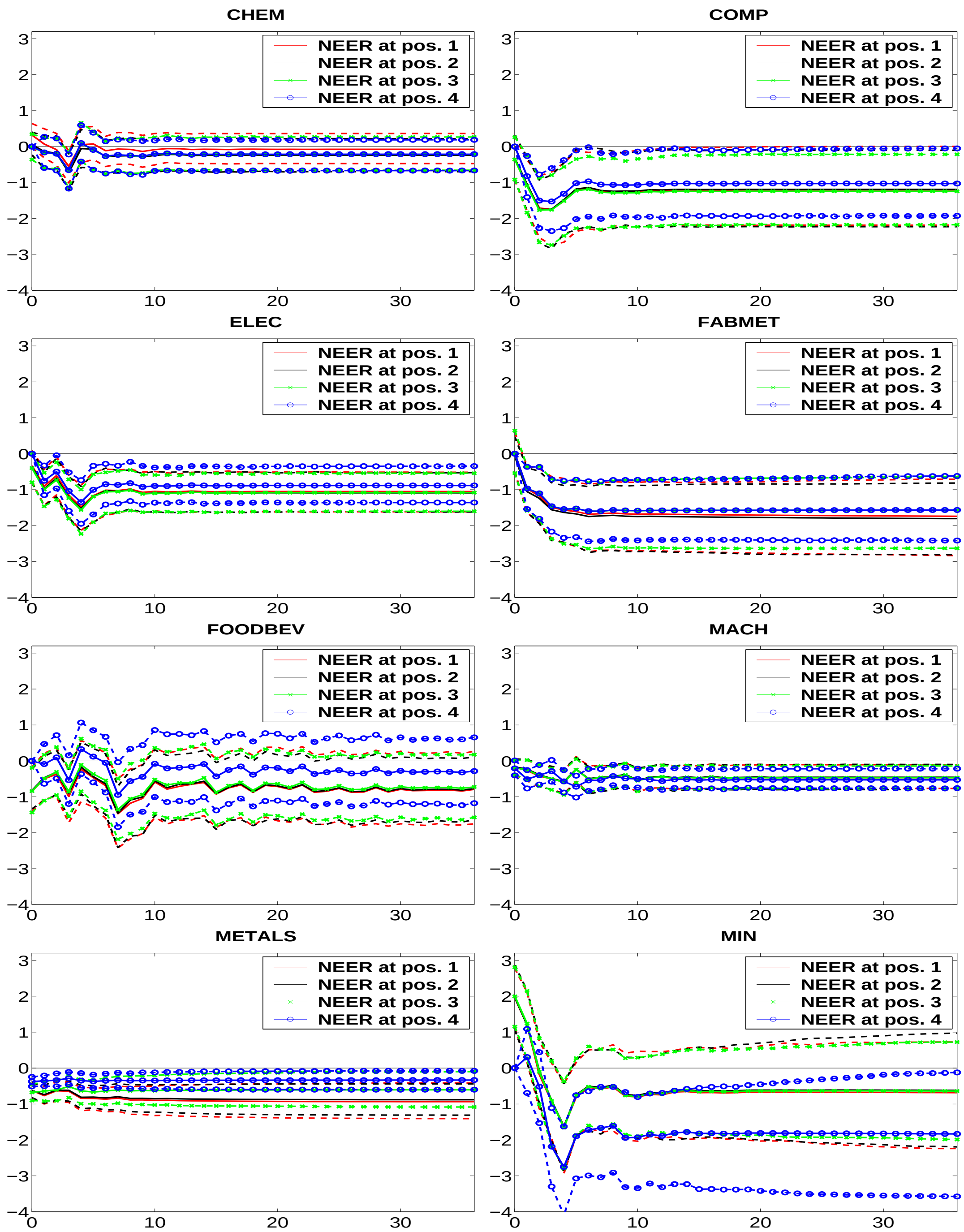

Figure 11: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$. The figures display the results for all positions of $\Delta neer$ in the Cholesky chain, see the main text and Table 6.

The solid lines display the point estimates and the corresponding dashed lines display 95% bootstrap confidence bounds.

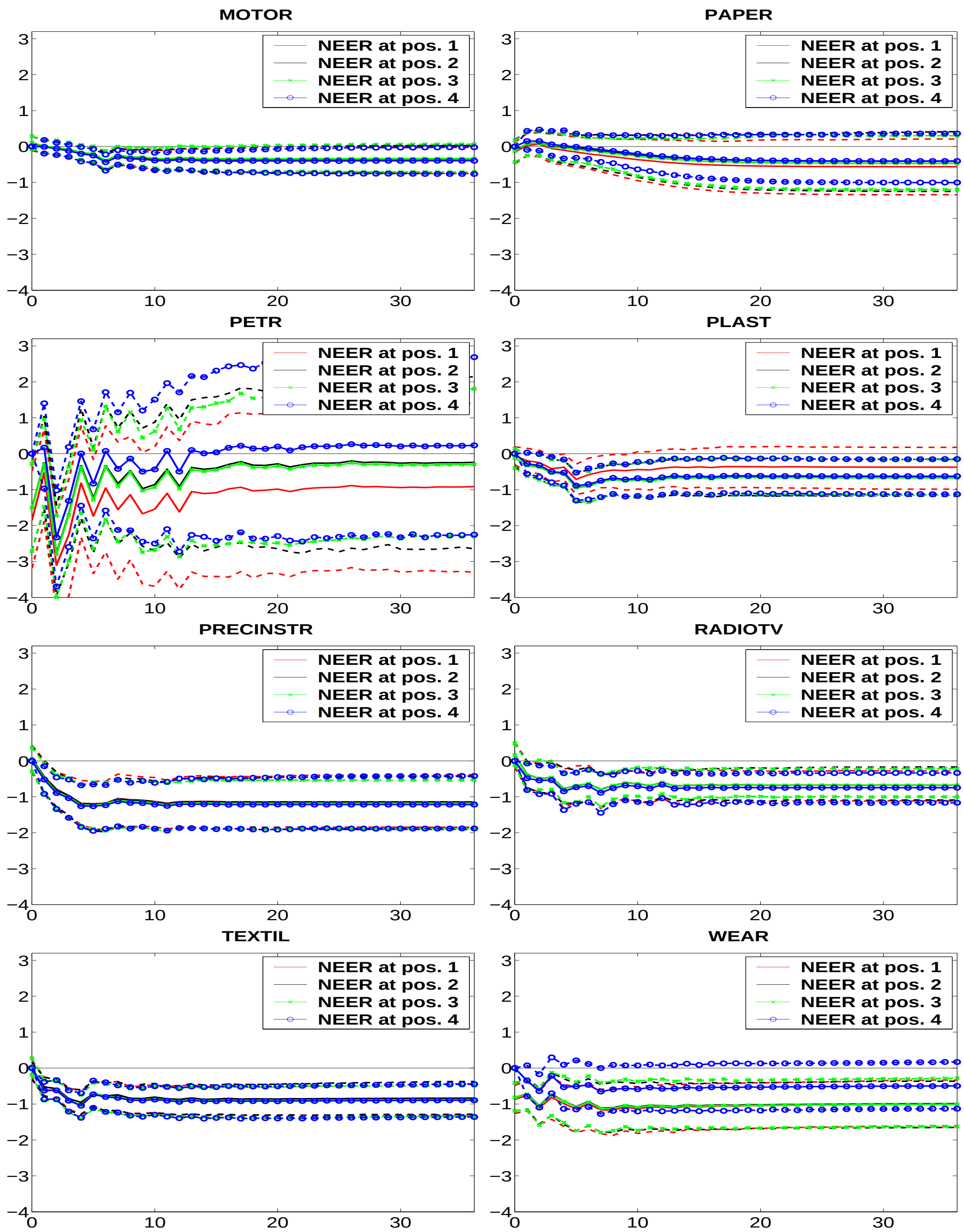

Figure 12: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$. The figures display the results for all positions of $\Delta neer$ in the Cholesky chain, see the main text and Tables 6 and 7.

The solid lines display the point estimates and the corresponding dashed lines display 95% bootstrap confidence bounds.

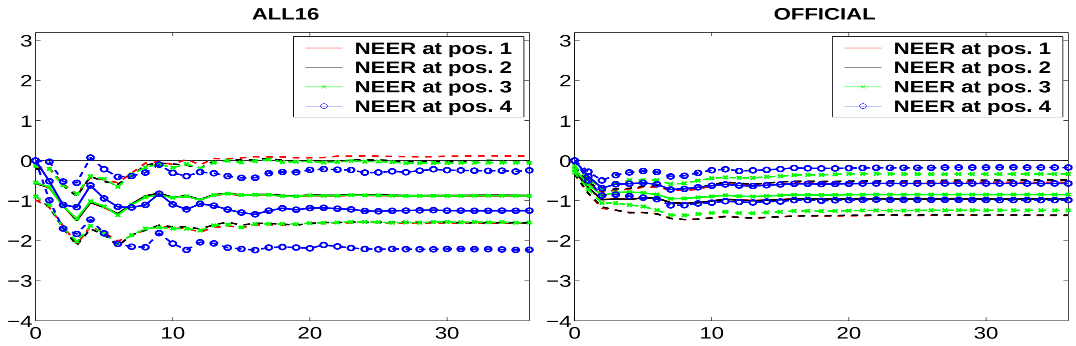

Figure 13: Accumulated impulse response functions for Δuvx from a unit shock to $\Delta neer$. The figures display the results for all positions of $\Delta neer$ in the Cholesky chain, see the main text and Table 7.
The solid lines display the point estimates and the corresponding dashed lines display 95% bootstrap confidence bounds.

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