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Abstract

This study investigates reactions to real exchange rate changes in the German, French and UK automobile and mechanical engineering sectors using monthly data from 1995 to 2010. Our findings indicate that EUR/USD appreciations hamper exports, but do not necessarily imply an aggravated business climate or export order-book assessment. This does not apply to the GBP/USD and corresponding time series for the UK. First and foremost, our fixed coefficient and time varying parameter VAR model estimates confirm the extraordinary role of the German key sectors, while currency union membership seems to play a minor role at best. Overall, the exchange rate susceptibility is less profound than claimed by lobbies and held as popular belief.

Keywords: Exporting sectors, confidence indicators, structural VAR

JEL classification: C30, E42, F41

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"The Euro should not fluctuate according to the mood of the markets. A monetary zone must have an exchange rate policy ... If not, we are insisting on countries making efforts to be competitive which are destroyed by the rising value of the Euro."

Francois Hollande (February 2013)

1   Introduction

Throughout the post-war period it is in particular the German economy that is widely regarded as depending on its exports. Public opinion sees exports as driving German –and ultimately also European– economic growth. Following this line of reasoning, key sectors such as the automobile and mechanical engineering industries are perceived as being especially susceptible to an appreciation of the Euro. On the other hand, they arguably profit from a depreciating Euro. In this context, the EUR/USD exchange rate is one of the most intensely observed relationships. In 2003, for example, the US business of German companies is said to have been decreasing by ten percent due to a weak Dollar (Belke *et al.* 2013). As of the second half of the first decade of the 2000s it became popular to talk about a "pain threshold" for European companies with regard to the EUR/USD exchange rate. Prominently the term is used referring to a suffering of German and/or French export industries from a strengthening of the Euro beyond a certain threshold. Recently, even terms such as "strategic depreciation" and "currency wars" are used to refer to the fear of European industries to lose grounds in competitiveness against the backdrop of devaluating competitor countries; see the above quote of French president Francois Hollande. However, with regard to stock market returns Griffin and Stulz (2001), for example, find that common shocks to industries across countries are substantially more important than competitive shocks due to changes in exchange rates. According to their findings, weekly exchange rate shocks explain but a tiny fraction of the relative performance of industries and, in particular, also of sectors that produce internationally traded goods.

In our reading of the existing literature the question of the existence and dimension of the exchange rate susceptibility of European core industries has not been satisfyingly answered. Several studies investigate the EUR/USD exchange rate dynamics in general at the level of a nation or supranation. See, among others, Sinn and Westermann (2001), Fratzscher (2008), and Belke *et al.* (2013). Another strand of literature analyzes the

issue of exchange rate susceptibility at the level of the firm or industry. Griffin and Stulz (2001), for example, investigate different sectors in the US, Canada, the UK, France, Germany, and Japan. As mentioned above, their results suggest that exchange rates play, if at all, only a minor role in determining sectoral performance. Contrary to Griffin and Stulz (2001), Bodnar and Gentry (1993) find significant exposure to exchange rate movements for industries in the US, Canada, and Japan. Williamson (2001) is a study that explicitly investigates the automobile industry finding, in general, significant exchange rate exposure for this sector. However, due to "lack of competition" Williamson excludes the German automobile industry from his analysis. Finally, Greenaway *et al.* (2010) emphasize the effects of exchange rate changes at a firm's cost side by noting that appreciation may have offsetting effects through relatively cheaper imported intermediate goods. Their empirical results for the UK manufacturing sector indicate that exchange rate dynamics might not have a significant effect at all once effects at the cost side are considered. Overall, there is no coherent picture of the EUR/USD exchange rate susceptibility of key exporting sectors of the core EU economies in the literature to back up the vast narrative evidence on the issue.

Here, we take a different route and define susceptibility more directly as an aggravation of both the sectoral order book indicators and of the business climate in the automobile and mechanical engineering industries. These two sectors account, for example, in the case of Germany for about one third of its total exports in 2011. Our central empirical strategy consists of estimating trivariate VAR models incorporating the real exchange rate, the volume index of exports from the two key exporting sectors, and sectoral confidence indicators using monthly data from 1995 to 2010. Against this backdrop, our study seeks to analyze three fundamental hypotheses:

- The popular belief of a profound exchange rate susceptibility of the core sectors, i.e. the automobile and mechanical engineering industries. [*H1*]

- The Williamson-hypothesis according to which the German economy and its key sector, the automobile industry, stand out (Williamson 2001). [*H2*]

- The Frankel-Rose-Melitz-hypothesis according to which currency union members get immunized as the union fosters intra-Eurozone trade counteracting exchange rate exposure (Frankel and Rose 2002, Melitz 2003). [*H3*]

To address the first point we focus on the German economy as our benchmark case.

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Comparing our results to estimates obtained from French sectoral series then allows us to assess a possibly exceptional reaction of German key sectors. Finally, considering the GBP/USD exchange rate along with UK sectoral series we analyze whether the two currency union members witnessed some change in their response to exchange rate shocks *over time*. Overall, our focus is on (foreign) demand-side effects of exchange rates, considering both realized demand, i.e. exports, as well as expected demand, i.e. export order-book assessments, and business confidence.

The remainder of the paper is organized as follows: Section 2 discusses sources of imperfect pass-through in the European context. In section 3, we use German series to estimate a reference case. Section 4 extends our reference case to French and UK time series. In section 5 we look at the sectoral response to an exchange rate shock over time. Finally, section 6 concludes.

2 Sources of Imperfect Pass-through

In a stylized model-world the reaction of exports and ultimately also of the business climate and other confidence indicators to changes in the real exchange rate in the exporting sectors is clear-cut. Both, in traditional Mundell-Fleming-Dornbusch type models (Dornbusch 1976) and in more recent open-economy models (Svenson 2000, Lane 2001) an appreciation (a depreciation) decreases (increases) the demand for domestically produced export goods by increaseing (reducing) their relative price. However, in a world of differentiated export goods supplied by firms producing with imported intermediate goods on segmented markets, characterized by imperfect competition and different shades of exposure to trade, the unambiguousness of this relationship gets lost. We will discuss four central sources of imperfect pass-through in the following: imported intermediate goods relating to [H1], imperfect competition relating to [H2], strategic firm behavior relating to [H1] and [H2], and exposure to trade relating to [H3].

A straightforward implication of an an appreciation of the domestic currency for firms producing with imported intermediate goods is reduced cost of such inputs. This can be seen as particularly important for European companies when it comes to paying the bill for energy used in the production of automobiles and machinery as neither the UK (making less than 3 percent of the world's oil production through North sea oil drilling over our sample period) nor Germany and France are among the major oil producing

economies in the world. Ultimately, this might fully counteract the pass-through for these industries (Greenaway *et al.* 2010).

In general, a full pass-trough requires perfectly competitive markets. Incomplete competition implies incomplete and inert price adjustments. For less than fully competitive markets a pass-through will typically be incomplete, referred to as pricing-to-market (PTM) in the literature (Krugman 1986). Accordingly, companies set their export price to $P = (1 + \pi)MC$, i.e., due to some market power they mark-up price their products.[1] PTM can be a possibility to preserve foreign market shares as it allows the exporter to stabilize supply prices in foreign currency by reacting to exchange rate movements with changes in the mark-up. PTM may also help to avoid menu cost in case of exchange rate changes which are only transitory in nature. Additionally, it might serve as a "war chest," whenever firms believe that regaining forgone market shares is more costly than transitory losses from foregone profits. Several empirical studies have found evidence for PTM on the side of exporters in general (Knetter 1993, Falk and Falk 2000) and for companies in the automobile and mechanical-engineering industry in particular (Belke *et al.* 2013). Ultimately, the extent of exchange rate pass-through and PTM is an empirical question hinged on the price elasticity of exports. PTM may help to cushion negative externalities of an appreciation on the demand side, although it will always go at the cost of firms' profits. Certainly, the extent of the latter depends on the price elasticity of exports. For example, according to Deutsche Bundesbank (2008), German exports, in particular, to non-European economies, react rather weakly to exchange rate changes.[2] If, for instance, the domestic currency appreciates by 1% real exports are estimated to fall by only about 0.25%. This is often ascribed to relative price-inelastic goods, for example, individualized investment goods such as specific machinery, making up a considerable share of German exports. Foreign customers may also stick to their supplier even though the value of foreign currency increases because of switching cost being even higher. As German companies are highly specialized in certain industries, "foreign consumers are 'caught' in their relation to German suppliers" (Belke *et al.* 2013).[3]

An alternative to cope with exchange rate risk from the firm perspective is to shift production abroad. This so called "natural hedging" has been widely practiced by the

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[1]This mark-up (price discrimination) may vary among export destinations.

[2]For the following arguments and figures see Belke *et al.* (2013).

[3]However, Chen *et al.* (2012) recently find no evidence that German exporters, in general, benefit from a demand elasticity advantage over other Eurozone countries.

German automobile industry in the past 15 years setting up production plants, particularly, in the US (e.g. BMW in Spartanburg, SC, in the US), but also recently in China (e.g. Volkswagen in Changchun and BMW in Shenyang). Though to a somewhat lesser extent, we also observe this behavior in the French automobile and machinery sectors. It might also motivate different forms of cooperations and mergers and acquisitions (e.g. PSA/General Motors, Daimler/Chrysler or ZF Friedrichshafen/TRW Automotive auto parts suppliers) along the value added chain. Natural hedging assures that local buyers are supplied avoiding exchange rate risks. An additional advantage lies in the possibility of firms to compensate losses from lowered exports with cheaper imports from the foreign country within their company as the domestic currency appreciates.

Exposure to trade is also seen and rationalized in the literature (Melitz 2003) as source of an imperfect pass-through. In our context, it particularly concerns the importance of intra-European trade. When investigating the EUR/USD exchange rate effect on the French or German business climate, one has to consider the role of other trading partners, in particular, intra-EU partners. The overall impact of a weak USD might be small, if forgone exports to the US can be compensated by trade with other (intra-EU) economies. Actually, the US is, for example, ranked third among the top trading partners of Germany in 2009. The list is headed by Europe and Asia: In 2009, 62% of the German exports went to EU member states, of which 17 are also members in the common currency area, while many of the remaining ones have pegged their currencies to the Euro. The latter concerns in particular the Central and Eastern European trading partners. Of course, this bears the implication for Germany that most of its production and trade becomes independent of exchange rate changes. According to estimates reported in Frankel and Rose (2002), a currency union roughly triples trade with other union members. As this externality is reasonably not realized over night but rather following a gradual process, it calls for a time varying analysis. In the sectoral model by Melitz (2003), which incorporates firm-level heterogeneity and counteracts pass-through via export market entry costs and intra-industry allocations across firms, only more efficient firms reap benefits from trade in the form of gains in market share and profit. Less efficient ones lose both. An increase in exposure to trade, according to Frankel and Rose (2002) implied by the installation of the Eurozone, hence, forces the least efficient firms out of the industry. As a consequence fostered intra-zone trade channels into productivity gains that may offset the pass-through of exchange rate shocks. Some micro-level evidence for this channel of transmission is found by Aw *et al.* (2000) and Pavcnik (2002) relying on East Asian and

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Latin American data, respectively.[4]

3 Empirical Assessment: the Case of Germany

3.1 Data

Our baseline empirical analysis relies on a sample consisting of monthly data of the EUR/USD real exchange rate (EXR), German exports, measured as volume index of exports (EXP), and the business climate (BC) in the relevant sectors: the automobile industry and the mechanical engineering sector. The period of observation ranges from January 1995 to October 2010 covering 190 observations. Data on the consumer price index (CPI) deflated EUR/USD exchange rate are taken from the Pacific Exchange Rate Service provided by the University of British Columbia (http://fx.sauder.ubc.ca/data.html). The exchange rate is denoted in price notation, i.e., $xEUR/1USD$. Hence, an increase of this ratio represents a depreciation of the Euro. Pre-1999 data for the Euro are the official ECU basket rates rather than imputed pseudo rates. Export series of the German automobile and mechanical engineering industry are obtained from Eurostat. The data cover all German exports within the SITC (*Standard International Trade Classification*) system product-group 7 "Machinery and Transport Equipment" to the United States. Exports are measured in 100 kilograms. The series is normalized to January 1995 (= 100). Our series of the business climate in the automobile and mechanical engineering sector are drawn from the detailed analysis of the "*ifo-Geschäftsklimaindex,*" which is published on a monthly base by the ifo Institute. The index is calculated as the mean of balances of percentage shares of positive and negative judgements reported by the companies with regard to their current business situation and their business expectations in the following six months (CESifo 2010, p. 3). Thus, as many other indicators, BC is based on two variables (business situation and business expectations), which are measured on a three-point Likert scale capturing a good, equal or bad state. For the period from 2004 to 2010 the business climate is reported for unified Germany for the automobile industry and the mechanical engineering industry, respectively. We calculate

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[4]Recently, Wierts *et al.* (2014) find that the sectoral composition of exports matters for the pass-through: Other things –in particular, characteristics of trading partners– equal, the pass-through is the more imperfect the higher the share of high technology exports. According to the OECD, motor vehicles as well as machinery and equipment are "medium-high technology" industries (Wierts *et al.* 2014, p. 939).

a joint business climate for the two sectors by averaging. Time series along with summary statistics are shown in Appendix A.1.

To judge whether one or all of the three series had to be seasonally adjusted, we used the variate difference method by Tintner *et al.* (1978). We find that our exports series (EXP) need to be seasonally adjusted.[5] Seasonal adjustment of the EXP series is done using standard US Census Bureau X-12 ARIMA. In a next step, we perform ADF tests for all series. We set and determine the maximum number of lags $k_{max} = \left[12(T/100)^{1/4}\right]$, where T denotes number of observations; here, $k_{max} = 14$. The optimal number of lags is chosen on the base of the AIC. The null of a unit root cannot be rejected at the 5%-level of significance for the series EXR and EXP (p-value equals 0.5927 and 0.5311, respectively).[6] We apply a Hodrick-Prescott filter (HP) with a smoothing weight of $\lambda = 129,600$ (Ravn and Uhlig 2002) to both series and consider log first differenced (logD filtered) series alternatively. Some reduced form analysis of the correlation structure of series can be found in Appendix A.2.

3.2   VAR model: estimation and model checking

This section describes our selection and estimation of a VAR[p]-model in reduced form

$$\boldsymbol{y}_t = \boldsymbol{c} + \boldsymbol{B}_1\boldsymbol{y}_{t-1} + ... + \boldsymbol{B}_p\boldsymbol{y}_{t-p} + \boldsymbol{u}_t \quad t = 1, ..., T, \tag{1}$$

where $\boldsymbol{y}_t = [y_{1t}, ..., y_{kt}]'$ denotes a $k \times 1$ vector containing observed series (EXR, EXP, BC; i.e $k = 3$), $\boldsymbol{B}_i$, $i = 1, ..., p$, are (fixed) $k \times k$ coefficient matrices, $\boldsymbol{c} = [c_1, ..., c_k]'$ denotes a (fixed) $k \times 1$ vector of intercept terms and $\boldsymbol{u_t} = [u_{1t}, ..., u_{kt}]'$ a k-dimensional *white noise* process with $E[\boldsymbol{u_t}] = \boldsymbol{0}$, $E[\boldsymbol{u_t}\boldsymbol{u_t'}] = \boldsymbol{\Omega}_u$ and $E[\boldsymbol{u_s}\boldsymbol{u_t'}] = \boldsymbol{0}$ for $s \neq t$.

The model order is set to $p = 3$. It is chosen on the base of the AIC and FPE information criterion ($p_{max} = 24$). Although, both criteria, the AIC as well as the FPE, overestimate the true order with positive probability (Lütkepohl 2006, p. 150), this choice seems to be unproblematic regarding degrees of freedom as well as estimation precision.[7]

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[5]Data on the business climate have been seasonally adjusted by the ifo Institute.

[6]The null of a unit root can be rejected at a 1%-level of significance for BC (corresponding p-value = 0.0002). Results are robust for different UR/stationarity tests such as the KPSS test.

[7]The negative effect of p being asymptotically too high can be corrected by estimating a subset VAR.

Stacking right hand side terms, we get

$$\boldsymbol{y}_t = \boldsymbol{X}_t' \boldsymbol{B} + \boldsymbol{u}_t \quad t = 1, ..., T, \tag{2}$$

where $\boldsymbol{B} = vec(\boldsymbol{B}_1, ..., \boldsymbol{B}_p)$ and $\boldsymbol{X}_t' = \boldsymbol{I}_k \otimes [1, \boldsymbol{y}_{t-1}', \boldsymbol{y}_{t-2}', ..., \boldsymbol{y}_{t-p}']$. We proceed with estimating the VAR[3]-model resorting to standard OLS. Results are shown in Table 4 and 4 of Appendix A.5. There is a number of insignificant coefficients. Following the principle of parsimony, we make parameter restrictions setting all insignificant parameters equal to zero. Estimating a subset VAR seems justified for two reasons. First, we know that p selected on the base of AIC and FPE is too large with positive probability implying, at least, the elimination of some lags. Secondly, putting zero constraints on insignificant parameters will improve the model's forecast precision which has a positive impact on the results of impulse response (IR) analysis and forecast error variance decomposition (FEVD) as both rely on estimated "quantities" (Lütkepohl 2006, p. 207; Lütkepohl and Krätzig 2004, p. 180). Estimates for the subset VAR are shown in Table 6 of Appendix A.5. Lagged values of EXR have a significant positive impact on EXP as well as BC. Obviously, this stands in contrast to the exchange rate being best explained exclusively by its own past values. For all of the three single equations we observe satisfactory high values for adjusted R-squares and F-test statistics.

The subset VAR[3]-model is stable if all eigenvalues of $\mathbf{B}$ have moduli smaller than one, i.e., $\det(\mathbf{I} - \mathbf{B}z) \neq 0$ for $|z| \leq 1$. This is equivalent to saying that the characteristic polynomial $\mathbf{B}(z) = \det(\mathbf{I} - \mathbf{B}_1 z - ... - \mathbf{B}_p z^p)$ of the VAR-filter $\mathbf{B}(L)$ has no roots in and on the complex unit root circle (Lütkepohl 2006). For our model, all eigenvalues of moduli of $\mathbf{B}$ are smaller than one (0.902, 0.857, 0.857, 0.404, 0.404, 0.390, 0.385, 0.271, 0.000). Formal tests for residual autocorrelation do not give rise to concerns. The multivariate Portmanteau test finds that the null $\Gamma_{\widehat{\mathbf{u}}}(\tau) = 0$ cannot be rejected at a 5%-level (p-value = 0.077),[8] i.e. $\widehat{\boldsymbol{u}}_t$ can be viewed as typical realizations of an uncorrelated process $\{\boldsymbol{u}_t\}$. This finding can be supported by a Breusch-Godfrey test that fits a VMA process

$$\widehat{\boldsymbol{u}}_t = \boldsymbol{B}_1 \widehat{\boldsymbol{u}}_{t-1} + ... + \boldsymbol{B}_q \widehat{\boldsymbol{u}}_{t-q} + \boldsymbol{\eta}_t, \tag{3}$$

where $\boldsymbol{\eta}_t \sim WN(\mathbf{0}, \boldsymbol{\Sigma})$. Following Lütkepohl and Krätzig (2004, p. 127) the number of lags for the Breusch-Godfrey statistics is smaller and equals 12 (p-value = 0.0680).

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[8]Lags used for the Portmanteau statistic = 30.

Performing a multivariate Jarque-Bera test comes to the conclusion that the residuals are non-normally distributed. More precisely, only for the residuals of the exchange rate equation the normal distribution cannot be rejected (p-value = 0.2228). As a consequence, forecasting intervals may not be reliable (Lütkepohl 2006, p. 174). However, in sum, we may conclude that the subset VAR[3] is an adequate approximation of the underlying data generating process. In line with basic model logics, we find a positive (negative) effect of a Euro depreciation (appreciation) on exports and business climate.

3.3   Identification and structural interpretation

In the following, we focus on the causal interpretation of the dynamics incorporated in the VAR-model. As we have seen above our subset VAR[3]-model is stable, guaranteeing that it has a VMA[∞]-represantation from which impulse response functions (IRF) can be calculated.

In disentangling the reduced form errors $\boldsymbol{u}_t$, which are correlated by definition, we –at first– rely on a Cholesky decomposition of the variance-covariance matrix $\boldsymbol{\Sigma}_u$. Orthogonalized shocks are given by $\boldsymbol{\epsilon}_t = C^{-1}\boldsymbol{u}_t$, where C is lower-triangular such that $CC' = \boldsymbol{\Sigma}_u$.[9] Together with our ordering of variables in $\boldsymbol{y}_t = [EXR_t, EXP_t, BC_t]'$ this decomposition constitutes a Wold causal chain running from the exchange rate over exports to the business climate. Note, in the context of our basic identification scheme it is crucial to stress that the only real macro-quantity is EXR, whilst all other variables in the system, including exports, represent sectoral quantities referring exclusively to the automobile and mechanical engineering industry. Hence, we implicitly allow only the non-sectoral variable to impact upon all other variables in t. The implicit assumtpion is that the dynamics of the EUR/USD exchange rate are exogenously determined. This assumption seems justified as the EUR/USD exchange rate equals the relative price of the domestic currency. This price, however, is determined by the amount of Euros supplied by the ECB which can be assumed to act independently of German sectoral exports and/or confidence. Clostermann and Schnatz (2000) identify factors that determine the real

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[9]Though model residuals are merely correlated, orthogonalized shocks make sense as they render impulse responses comparable. However, in case IRFs (or the FEVDs) are based on orthogonalized shocks results might be sensitive with regard to the ordering of variables. In our example this objection can, at least partly, be rejected by referring to the merely present instantaneous residual correlation (Lütkepohl and Krätzig 2004, p. 181). For further information on the robustness of results with regard to different variable orderings see Section 3.4.

EUR/USD exchange rate. According to their study, there is no need to assume that the exchange rate should be instantaneously influenced by German car and/or machinery exports to the US or by the business climate in the respective industries.[10] Our assumption can also be supported in a more technical way by testing for Granger-causality. The hypothesis of "No instantaneous causality between: EXR and EXP, BC" can not be rejected at a 5% level of significance (p-value $= 0.2151$). German sectoral exports on the other hand are assumed to be contemporaneously influenced by the exchange rate. The assumption that the exchange rate in t impacts on exports in t does not seem to be problematic either. There are several (empirical) studies that justify our belief in such a contemporaneous relation (e.g. De Grauwe and Verfaille 1988, Asseery and Peel 1991, Sauer and Bohara 2001). It is also reasonable to presume –as we do through our ordering– that sectoral export volumes are not contemporaneously influenced by sectoral confidence indicators, while the industries' business climate might be contemporaneously influenced by sectoral export activity.

Finally, our specification assumes that the business climate is instantaneously influenced by the exchange rate and exports but not vice versa. This is a plausible assumption with regard to the construction of the used indicators. As mentioned earlier, the business climate index published by the *ifo* institute is calculated as the mean of balances of percentage shares of positive and negative judgements reported by companies with regard to their current business situation and business expectations for the following six months. The current business situation and business expectations are measured on a 3-level-Likert scale representing a state $s \in S = \{+, =, -\}$. We may think of a variable on such a scale as resulting from an unobserved process $f(t)$. This idea can be formalized as follows: $s_{i,t}^* = f(t) + \nu_t$ with $\nu_t \sim N(0, \sigma^2)$, $i = 1, ..., n$ and $t = 1, ..., T$. However, as survey participant i must answer on the categorial scale we only observe

$$
s_{i,t} = \begin{cases} + \text{ for } s_{i,t}^* > \tau^+ \\ = \text{ for } \tau^+ \geq s_{i,t}^* > \tau^- \\ - \text{ for } \tau^- \geq s_{i,t}^*. \end{cases}
$$

While the exact shape of $f(t)$ might be undefined, it is usually assumed that the unobserved process represents the business cycle (Seiler 2012). Yet, it seems not unrealistic to

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[10] Of course, German exports might influence the EUR/USD exchange rate via a change in domestic national income. However, the necessary adjustment processes will take some time.

assume that $f(t)$ by anticipation of the survey participants also captures to some extent assessments of exchange rate dynamics and/or the development of exports to trading partners.

Due to these considerations we are confident that the Cholesky decomposition along with the chosen ordering of variables in $\boldsymbol{y}_t$ supplies us with structural exchange rate shocks that may be used for causal interpretation in a depreciation scenario.
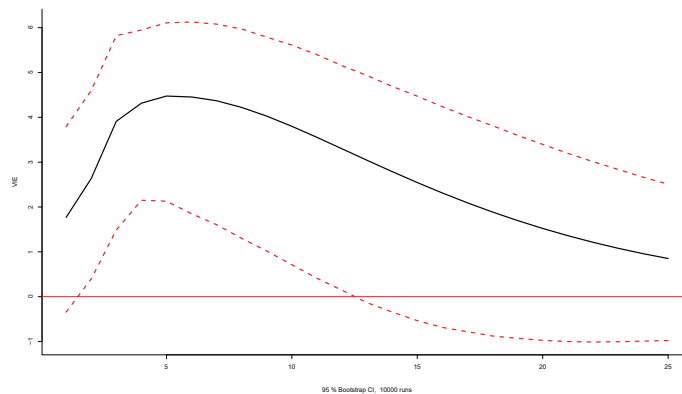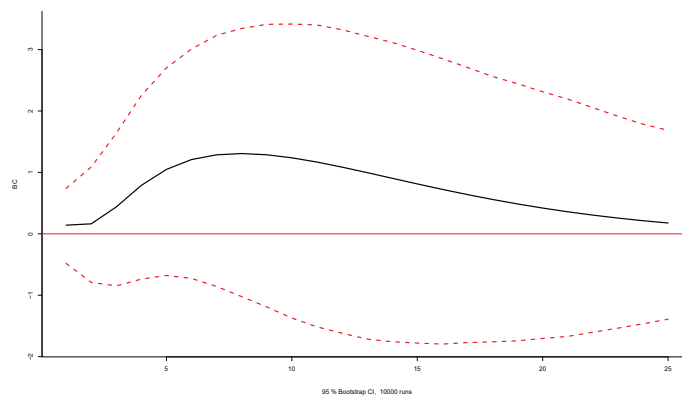

Figure 1. Response of EXP to EXR shock: Germany

Figure 2. Response of BC to EXR shock: Germany

Our strategy is to focus on the response of exports and business climate to orthogonalized shocks from the exchange rate. Corresponding IRFs are plotted in Figures 1 to 4. Overall, it seems that exports as well as the business climate react positively to a shock in the exchange rate. Exports show a rather instantaneous and significantly positive response in the short run. After about thirteen periods exports are back to normal, i.e.

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the bootstrapped confidence intervals[11] cover the zero line again. Conclusions are quite different regarding the reaction of the business climate. First of all, BC seems to react in a less clear-cut way to a shock in the exchange rate. Secondly, the response is delayed for about two periods. Thirdly and most importantly, the confidence intervals clearly include the zero line allowing us not to speak of a significantly positive response of the business climate. In other words, exchange rate changes can leave the business climate unchanged.
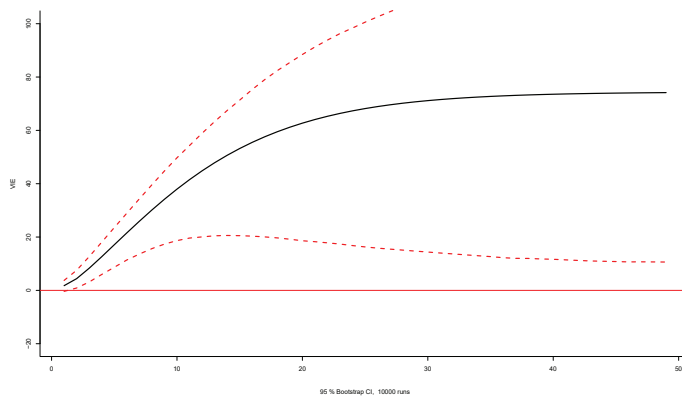
Figure 3. Response of EXP to cummulative EXR shock: Germany
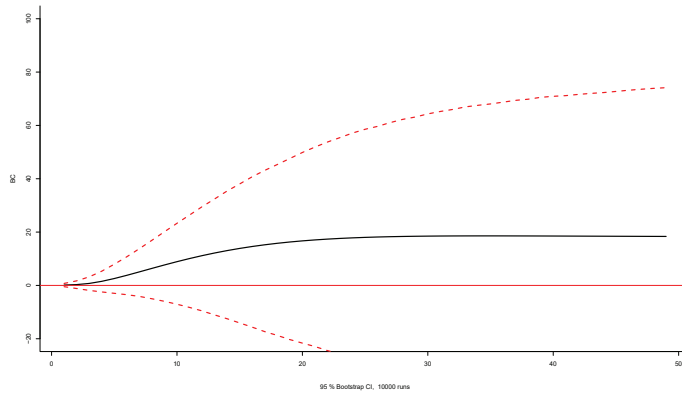
Figure 4. Response of BC to cummulative EXR shock: Germany

Figure 3 and Figure 4 also show the cumulated IRFs of EXP and BC which basically confirm our results. However, in a system of fully flexible exchange rates (free float), as in the case of the EUR/USD exchange rate, it would not be reasonable to assume

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[11]95% confidence intervals were bootstraped with 10,000 runs.

constant shocks, i.e. an ongoing depreciation or appreciation of the Euro.[12]

In a next and final step of our reference case, it is straight forward to look at the FEVD for the three variables. The forecast horizon is 12 months. Again by the variables ordering in the left hand side vector $\boldsymbol{y_t}$ the following causal structur is assumed: exchange rate $\rightarrow$ exports $\rightarrow$ business climate. Not surprisingly for the subset VAR[3] the error variance of the exchange rate's h-step forecast is exclusively explained by own innovations. This result does not fundamentally change, if one looks at the unrestricted VAR[3]: The maximum contribution of EXP (BC) to forecast error variance is 6.0% (3.8%). Looking at the FEVD of exports, we see that innovations in the exchange rate gain significant impact on the variable's dynamics. For $h = 12$ shocks of the exchange rate make up 30.9% of exports forecast error variance. Also the business climate has a strong effect on the dynamics of the EXP series (up to 28.5% for $h = 12$). Finally and surprisingly, we observe that the exchange rate makes little to no contribution to the dynamics of the business climate. At maximum 3.6% of the business climate's forecast error variance can be explained by innovations in the exchange rate. Of course, the exports' impact is somewhat stronger (amounting to up to 13.1%). Dynamics of the business climate indicator are best explained by own innovations. Results are summarized in Figure 5.
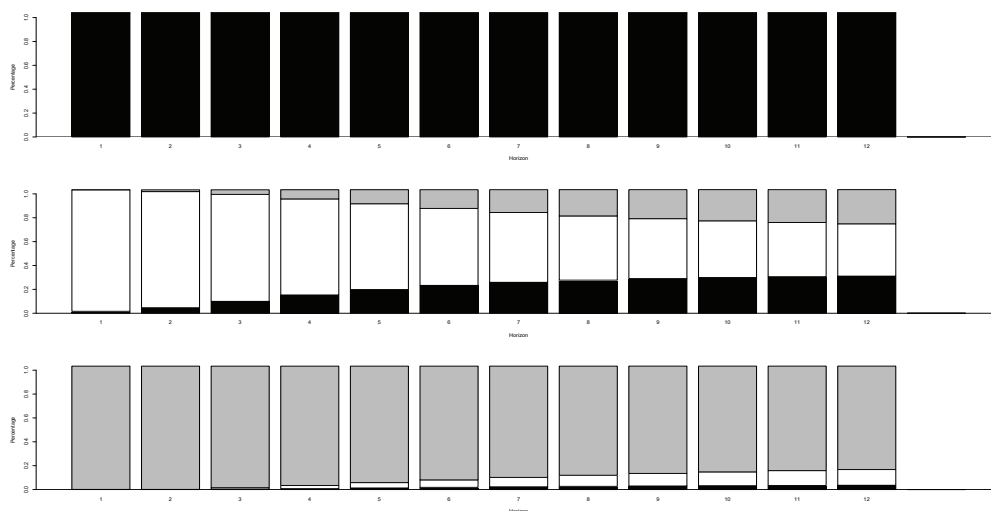
Figure 5. Forecast error variance decomposition (FEVD): Germany
Note: black shaded: EXR, white shaded: EXP, grey shaded: BC
FEVD from first to third row for EXR, EXP, BC

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[12]Nevertheless, it might be possible for fixed exchange rate regimes. A constant shock would correspond to a revaluation of the domestic currency.

3.4 Robustness of results

To ensure that our results are not artificially generated by the respective filtering method or climate indicator we consider both logD-filtered series and alternative BC indicators. Figures given in Appendix A.4 show the IRFs and FEVD for different filtering methods (HP and logD) and additional climate indicators for the respective industries (i.e. the Confidence Indicator and the Assessment of Export Order-Book Indicator published by the European Commission). Our results are robust with regard to different filtering methods. In particular the application of the logD filter does not qualitatively change our main finding of a significant reaction of exports and an insignificant reaction of the business climate to an exchange rate shock. Results for the logD-filtered series are shown in Figure 4 of Appendix A.4. The use of alternative climate indicators (i.e. the Confidence Indicator and the Assessment of Export Order-Book Indicator published by the European Commission) neither alters our results in a qualitatively relevant way as can be seen by comparing IRFs in Figure 3 and Figure 4 of Appendix A.4 column by column.

There might remain some objections regarding the use of a business climate or confidence index as proxy for "pain" anticipated by firms against the backdrop of a strengthening Euro. One could argue that exports play only a minor role with regard to firms' profits compared to domestic sales. If this holds true, an omission bias is a probable caveat. In this case the insignificant reaction of the climate indicator to an exchange rate shock would neither be unexpected nor would it be an indication for hedging strategies or the like. However, there are two arguments speaking against such objections: First, in our baseline analysis we focus on the German automobile and mechanical engineering industries. For our period of observation, cars and machines clearly represent Germany's most important export goods. Hence, exports most reasonably play a major role for these industries. Secondly, our results are robust with regard to the use of the Assessment of Export Order-Book Indicator. This ensures that only firms are surveyed that actually engage in exporting and that judge their profit situation against the backdrop of export activity. A positive shock to the exchange rate does also not imply a significant impact on the firms' assessment of their export order-book levels; see respective last column of Figure 3 and Figure 4 in Appendix A.4. This suggests that transitory exchange rate shocks are not internalized in order levels expectations of German companies in the relevant sectors.

Although the ordering of variables in $\boldsymbol{y}_t$ is reasonable, we also follow the suggestion by Sims (1981) and try various triangular orthogonalizations. It can be shown that our main result of a significant reaction of exports and an insignificant reaction of the respective climate indicator to an exchange rate shock is not altered by different variable orderings.

Notice that all our results do not change qualitatively if we shorten our sample, i.e. either focus on the post-1998 period or leave out the 2008/09 financial crisis.

Finally, we also checked the robustness of our results with regard to the use of trade-wheighted exchange rates instead of bilateral exchange rates, i.e. we used the real effective exchange rate (CPI deflated, Euro area-17 countries vis-à-vis the EER-12 group of trading partners: AU, CA, DK, HK, JP, NO, SG, KR, SE, CH, GB and US), which is supplied by the Deutsche Bundesbank. Again, this leaves our results qualitatively unchanged. Detailed findings are available on request from the authors.

4   France and the United Kingdom

We have shown that our main finding using German data is robust with regard to a range of different model specifications. However, do these results apply to other countries' core sectors or do they rather confirm the Williamson-hypothesis [H2] of a profound distinctiveness of the German automobile and machinery industry? In order to answer this question, we conduct the analogue analysis using data for France and the United Kingdom.[13] Results are summarized in Figures 7 to 14 of Appendix A.4. In the case of France, findings from the IR analysis look, at first sight, similar to the German case. However, while there is no significant pass-through of exchange rate shocks down to the confidence or order-book assessments in the case of HP-filtered series, the logD-filtered series generate a slim and weakly significant negative reaction of the French sectors' confidence and order-book indicators. As it represents a net effect, it can be either rooted at the cost side if French firms in the automobile and machinery sectors hinge on many US-imported intermediate goods or in an adverse effect on firms' balance sheet positions[14] or in a mixture of both outweighing the perceived positive effect on exports.

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[13]As climate indicator for the respective industries the European Commission Confidence and Assessment of Export Order-Book Indicator were used again.

[14]A dollarized liabilities postion (i.e. debt service predominantly denominated in US dollars and revenues predominantly realized in domestic currency) deteriorates balance sheets and,

This finding of a fragile reaction depending on different filtering devices carries over to the results of the FEVD. While in case of the HP-filtered series the exchange rate has no significant impact on the climate indicators' dynamics, for the log first differenced series we observe that in case of the Confidence Indicator (Assessment of Export Order-Book Indicator) at maximum 11.7% (8.6%) of the forecast error variance is explained by exchange rate innovations. Compared to the German case this represents quite a share. Results are less ambiguous in case of relying on UK series. The positive reaction of exports is tiny and insigificant at conventional levels. Additionally, from the IRFs there is robust evidence that a positive shock to the GBP/USD exchange rate, i.e. a depreciation of the British Pound against the Dollar, has a significant and comparatively strong and deteriorating effect on both climate indicators. More than in the French case, it is reasonable and straightforward to ascribe the direction of this reaction to an outweighing negative balance sheet effect of depreciation and/or a deterioration at the cost side of firms due to dependence on US-imported inputs. In line with this finding, exchange rate innovations explain quite a substantial part of the indicators' forecast error variance, i.e. up to 16.3% (21.8%) in case of the (logD-filtered) Confidence Indicator (Assessment of Export Order-Book Indicator).

To sum up, in terms of exchange rate susceptibility or pass-through of exchange rate shocks down to the assessment of current and future demand we find the following qualitative ordering of national core sectors: UK (significant pass-through), France (shaky pass-through), Germany (no significant pass-through).

This ordering, at least, leaves open some possibility of a higher immunity against exchange rate changes of the German and French as opposed to the UK automobile and mechanical engineering industries which might be an indirect product of the currency union and the common market. In this sense and against the backdrop of the Optimum Currency Area (OCA) approach, ultimately representing a cost-benefit approach, the cushioning currency union effect that in the end lets exchange rate changes not undermine the general business outlook in the core sectors might even be a part of the "uncommon arguments for common currencies" (Mundell 1973). However, as the Eurozone was not installed over night but in phases and the immunization in the sense of the Frankel-Rose-Melitz-hypothesis [H3] sketched in Section 1 and Section 2 might develop over time, we

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in the presence of financial frictions, causes the external finance premium to increase and, consequently, investment and expected demand to decrease; see Aghion *et al.* (2000) and Bahmani-Oskooee and Miteza (2003).

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try to take account for these aspects in the following section.

5   The Role of the Currency Union

As argued above, if the Frankel-Rose-Melitz-hypothesis [*H3*] holds, the effect of an exchange rate shock on the business climate in the exporting industries should have changed over time. As trade with other currency union members gained in weight, vis-à-vis to the US, the effect of a Euro depreciation or appreciation on the business climate and export assessments might have decreased over time. Consequently, an extraordinarily strong DM, Franc, or ECU implied a strain on the business outlook of German and French exporters that was more severe before the Eurozone than nowadays. This is due to the fact that today foregone sales to the US, due to a strong Euro, can be more easily offset by trade with other Eurozone countries. Additionally, the Frankel-Rose-Melitz argument also applies to intermediate production leading to fostered intra-zone trade of intermediate goods counteracting a pass-through due to the dependence on high shares of US-imported intermediate goods. Both channels are supposed to not (fully) operate for countries that became not members of the currency union. Our period of investigation covers several crucial milestones and phases of Stage II (1994-1998) and Stage III (beginning 1999) of the EMU. For Stage II, we cover, in particular, the set-up of the ERM-II in mid-1997 and the creation of the ECB in mid-1998. For Stage III, i.e. the beginning of the single monetary policy regime of the ECB, we cover the three-year transition period starting 1999 before the Euro cash changeover in early 2002.[15]

In order to account for this development and to model indicator responses to an exchange rate shock over time we fit a T[ime]V[arying]P[arameter]-VAR with stochastic volatility à la Primiceri (2005) to our data. Note, although such a model does not require stationary series and, hence, does not require filtering of our series, we proceed in analogy to the preceding analysis. This strategy ensures comparablity of standard VAR and TVP-VAR based findings.

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[15]For evidence of effects of the changeover and inert adjustment to the new common currency on sentiment and confidence at the individual level see Wunder *et al.* (2008).

5.1 TVP-VAR model: estimation and model checking

We fit the following model to the data:

$$\boldsymbol{y}_t = \boldsymbol{c}_t + \boldsymbol{B}_{1,t}\boldsymbol{y}_{t-1} + ... + \boldsymbol{B}_{p,t}\boldsymbol{y}_{t-p} + \boldsymbol{u}_t \quad t = 1, ..., T, \tag{4}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector containing the (logD) exchange rate, exports, and the business climate, respectively, $\boldsymbol{c}_t$ is an $n \times 1$ vector of time-varying intercepts, $\boldsymbol{B}_{i,t}$, $i = 1, ..., p$ are $n \times n$ matrices of time-varying coefficients and $\boldsymbol{u}_t$ heteroskedastic unobservable shocks, where $\boldsymbol{u}_t \sim N(0, \boldsymbol{\Omega}_t)$. As the model is trivariate, $n = 3$. In terms of $\boldsymbol{\Omega}_t$ we can think of the following triangular reduction

$$\boldsymbol{A}_t \boldsymbol{\Omega}_t \boldsymbol{A}_t' = \boldsymbol{\Sigma}_t \boldsymbol{\Sigma}_t',$$

with $\boldsymbol{A}_t$ denoting the following lower triangular matrix

$$\boldsymbol{A}_t = \begin{bmatrix} 1 & 0 & 0 \\ \alpha_{21,t} & 1 & 0 \\ \alpha_{31,t} & \alpha_{32,t} & 1 \end{bmatrix}$$

and $\boldsymbol{\Sigma}_t$ being the diagonal matrix

$$\boldsymbol{\Sigma}_t = \begin{bmatrix} \sigma_{1,t} & 0 & 0 \\ 0 & \sigma_{2,t} & 0 \\ 0 & 0 & \sigma_{3,t} \end{bmatrix}.$$

From this we obtain

$$\boldsymbol{y}_t = \boldsymbol{c}_t + \boldsymbol{B}_{1,t}\boldsymbol{y}_{t-1} + ... + \boldsymbol{B}_{p,t}\boldsymbol{y}_{t-p} + \boldsymbol{A}_t^{-1}\boldsymbol{\Sigma}_t\boldsymbol{\epsilon}_t \quad t = 1, ..., T, \tag{5}$$

with $\boldsymbol{\epsilon}_t \sim N(0, \boldsymbol{I}_n)$. By $\boldsymbol{B}_t = vec(\boldsymbol{B}_{1,t}, \boldsymbol{B}_{2,t}, ..., \boldsymbol{B}_{p,t})$ equation (5) can be rewritten in the following way

$$\boldsymbol{y}_t = \boldsymbol{X}_t'\boldsymbol{B}_t + \boldsymbol{A}_t^{-1}\boldsymbol{\Sigma}_t\boldsymbol{\epsilon}_t \quad t = 1, ..., T \tag{6}$$

with

$$\boldsymbol{X}_t' = \boldsymbol{I}_n \otimes [1, \boldsymbol{y}_{t-1}', \boldsymbol{y}_{t-2}', ..., \boldsymbol{y}_{t-p}'],$$

where $\otimes$ denotes the Kronecker product. The model parameters evolve according to an AR(1) process

$$\boldsymbol{B}_t = \boldsymbol{B}_{t-1} + \boldsymbol{\nu}_t, \quad \boldsymbol{\nu}_t \sim N(0, Q), \tag{7}$$

$$\boldsymbol{\alpha}_t = \boldsymbol{\alpha}_{t-1} + \boldsymbol{\zeta}_t, \quad \boldsymbol{\zeta}_t \sim N(0, S), \tag{8}$$

$$log(\boldsymbol{\sigma}_t) = log(\boldsymbol{\sigma}_{t-1}) + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim N(0, W), \tag{9}$$

with $\boldsymbol{\alpha}_t$ being the non-zero and non-one element of matrix $\boldsymbol{A}_t$ and $\boldsymbol{\sigma}_t$ being the vector of the diagonal elements of $\boldsymbol{\Sigma}_t$.[16] S is assumed to be block-diagonal, i.e.

$$S = Var(\boldsymbol{\zeta}_t) = \begin{bmatrix} S_1 & 0_{1\times2} \\ 0_{2\times1} & S_2 \end{bmatrix},$$

where $S_1 = Var(\zeta_{21,t})$ and $S_2 = Var([\zeta_{31,t}, \zeta_{32,t}]')$, with $Var(\cdot)$ denoting the variance operator.

Together, equations (6) to (9) form a state-space representation of the TVP-VAR. The model innovations are assumed to be jointly normal. Again the model order has been chosen to be $p = 3$.

In estimating the model we follow a Bayesian approach rather than relying on Maximum Likelihood (ML) estimation. This is due to the fact that, given the high dimensionality and nonlinearity of the problem, ML-estimates are not efficient.[17] Even when being able to compute an ML-estimate of which one can be sure that it is not just a local maximum, it remains unsettled how to deal with the uncertainty related to the estimate. Bayesian inference deals with this problem by evaluating the posterior distributions of states and parameters and thus incorporates the uncertainty about these quantities. We will use a Gibbs sampler, which is a special variant of Markov Chain Monte Carlo (MCMC) methods, to determine the posterior distributions of $\boldsymbol{B}^T$, $\boldsymbol{A}^T$, $\boldsymbol{\Sigma}^T$, and the hyperparameters (Q, S, W).[18]

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[16]While $\boldsymbol{B}_t$ and $\boldsymbol{A}_t$ are modeled as random walks without drift the stochastic volatilities follow a geometric random walk. The random walk assumption for the coefficients and the contemporaneous relations comes along with a number of disadvantages; see Primiceri (2005) for a detailed discussion. However, it has the great advantage of reducing the number of parameters to be estimated.

[17]One particular drawback is that a complicated model as defined in the previous section will usually come along with a Likelihood function including multiple peaks; see Primiceri (2005) for a detailed discussion of problems related to ML-estimation of state space models.

[18]Superscript $(\cdot)^T$ indicates that all information up to point T is used in the estimation of

We assume that the initial states for $\boldsymbol{B}^T$, $\boldsymbol{A}^T$, $\boldsymbol{\Sigma}^T$ and the hyperparamteres are independent of each other. Furthermore, it is assumed that the priors $p(\boldsymbol{B}_0)$, $p(\boldsymbol{\alpha}_0)$ and $p(log(\boldsymbol{\sigma}_0))$ are normally distributed, whereas the priors for Q, W and S follow an independent inverse-Wishart distribution with scale matrix Ψ and degrees of freedom m. Estimating a time-invariant VAR by OLS on a small training sample of 30 observations is used to calibrate the priors. Letting $\hat{x}$ denote the point estimate of an unknown parameter x and $\hat{V}_{\hat{x}}$ the respective variance, we get the following priors

$$\boldsymbol{B}_0 \sim N(\hat{\boldsymbol{B}}_{OLS}, 3 \cdot \hat{V}_{\hat{\boldsymbol{B}}_{OLS}}), \tag{10}$$

$$\boldsymbol{A}_0 \sim N(\hat{\boldsymbol{A}}_{OLS}, 3 \cdot \hat{V}_{\hat{\boldsymbol{A}}_{OLS}}), \tag{11}$$

$$log(\boldsymbol{\sigma}_0) \sim N(log(\hat{\boldsymbol{\sigma}}_{OLS}), 3 \cdot \boldsymbol{I}_3), \tag{12}$$

$$Q \sim IW(k_Q^2 \cdot 30 \cdot \hat{V}_{\hat{\boldsymbol{B}}_{OLS}}, 30), \tag{13}$$

$$W \sim IW(k_W^2 \cdot 3 \cdot \boldsymbol{I}_3, 3), \tag{14}$$

$$S_1 \sim IW(k_S^2 \cdot 2 \cdot \hat{V}_{\hat{\boldsymbol{A}}_{1,OLS}}, 2), \tag{15}$$

$$S_2 \sim IW(k_S^2 \cdot 3 \cdot \hat{V}_{\hat{\boldsymbol{A}}_{2,OLS}}, 3), \tag{16}$$

with $k_Q = 0.01$, $k_S = 0.1$ and $k_W = 0.01$. For (W, S) the degrees of freedom m are set to one plus the dimension of the respective matrix. This is in any case the minimum number of degrees of freedom in order for the inverse-Wishart distribution to be properly specified (Primiceri 2005). In case of (Q) the degrees of freedom m are set to 30, which is the size of the training sample and leads to a slightly tighter prior. Except for W, for all the priors on the hyperparameters the scale matrices are constant fractions of the variances resulting from the model fitted to the training sample. Chosen in that way the priors are not completely uninformative yet still rather diffuse. This guarantees that the information embodied in the priors is soon dominated by the information contained in the data.

Simulating the joint posterior distribution $(\boldsymbol{B}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, S, W)$ via Gibbs sampling takes place according to the following steps: Sequentially draw $(\boldsymbol{B}^T)$, $(\boldsymbol{A}^T)$, $(\boldsymbol{\Sigma}^T)$ and the hyperparameters (Q, S, W) given the data and the rest of the parameters.[19] We perform 30,000 sampling iterations, discarding the first 20,000 draws as burn-in phase. The Gibbs sampler, as every Markov chain sampler, is a dependence chain algorithm,

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parameters of interest. This is because MCMC is a smoothing rather than a filtering method.

[19]See Appendix A.3 for a detailed description of the separate Gibbs sampling steps.

i.e. the different draws are not independent from each other. In order to break the correlation between the different draws, we keep only every tenth draw. Additionally, following Cogley and Sargent (2001), all draws for the coefficient vector that would lead to an explosive solution of the model are discarded.

5.2 The business climate response over time

Figures 6 to 8 display the absolute[20] response of the sectoral business climate indicator to an exchange rate shock after six months (6M) by the respective solid line country by country. Additionally, the running mean, i.e. a mean that is calculated anew with each new observation, of this response over time is shown by a dotted line in each diagram. For the sake of comparability ordinate values are identically scaled. Obviously, the medium-term business climate's reaction to an exchange rate shock developed rather differently for the different economies.

Responses of the German and UK automobile and mechanical engineering sectors conspicuously drift apart with the actual implementation of the single currency for the common market in January 2002.[21] The starkest contrast is given comparing the time-varying response function of German core exporting sectors (Figure 7) with the corresponding one for the UK (Figure 6). Although Germany starts from a slightly lower level of medium run exchange rate susceptibility of its core sectors, it clearly falls over nearly the entire period of observation. The opposite applies in the case of the UK. For French key exporting sectors the reaction of business confidence to exchange rate shocks has been rather stable over the total period of observation (Figure 8). The running mean (dotted line) is basically constant, suggesting no change in responsiveness over time apart from a slight increase coinciding with the begining of the global crisis. However, the running mean suggests this increase in exchange rate shock susceptibility to rather represent transient changes that are transitory in nature. An immunization in the sense of the Frankel-Rose-Melitz-hypothesis [H3] is not confirmed for French data. Hence, our TVP-VAR based findings leave the impression that it is not the coming into being of

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[20]From the constant parameter VAR impulse responses, we know that a pass-through might result in a significant positive or negative reaction of the expected demand indicator. As we are rather after tendencies or trends, we focus on absolute responses and abstract from calculating confidence bands from the posterior distributions.

[21]In case of France the response is plotted only from 1999 onwards as the training sample consists of 50 as opposed to 30 observations for the other countries.

the Eurozone membership that has led to a jointly declining responsiveness of Eurozone members' sectoral business confidence to a EUR/USD exchange rate shock. For though the German sectoral index reaction clearly appears to markedly decrease after the Euro cash changeover, we see no indication for such a tendency in the case of the French series.

First and foremost, our findings again make the point of an extraordinary role played by the German key sectors. It is most probably rooted in the high specifity and high price-inelasticity of their products as well as in the proliferation of natural hedging strategies. In line with our results reported in Section 3 and Section 4, we regard the Williamson-hypothesis [H2] as confirmed.
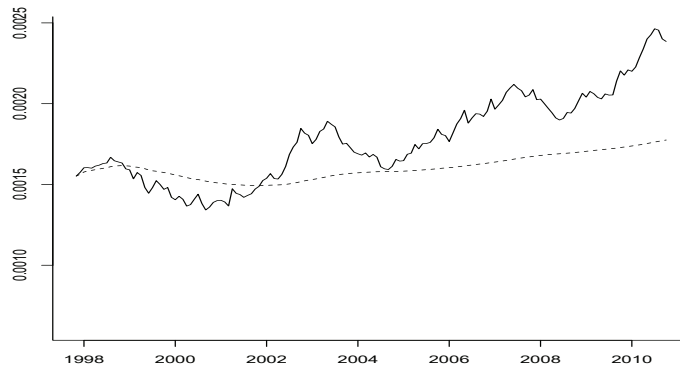
Figure 6. 6M response of BC to EXR shock: UK
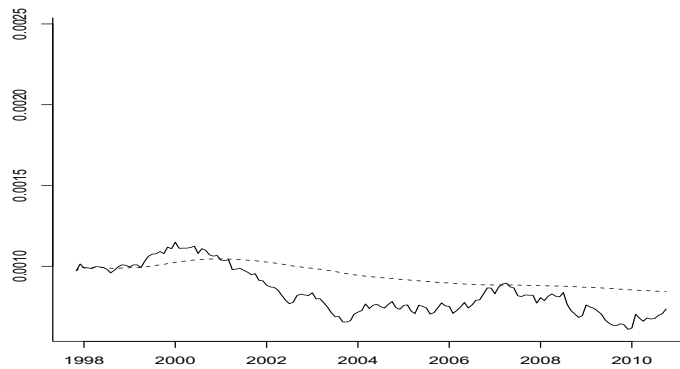
Figure 7. 6M response of BC to EXR shock: Germany
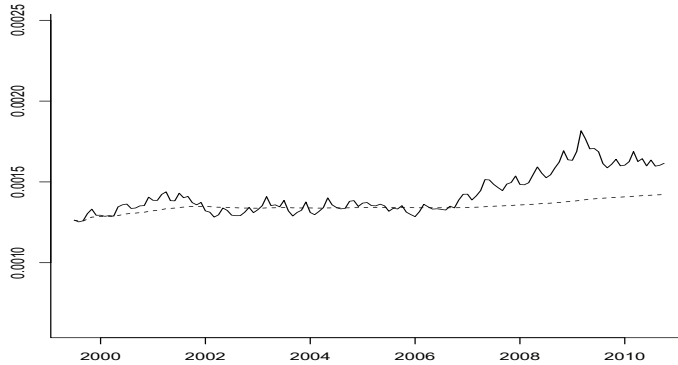
Figure 8. 6M response of BC to EXR shock: France

6 Interpretation and Policy Implications

The German automobile and related industries strategically use offshoring to the US quite differently than, for example, to China, where the prime target is to directly serve a strongly growing regional demand. According to Deutsche Bank Research (2014), virtually 100 percent of German vehicles produced in China are sold in China. The motivation for German offshore-plants in the US is quite different. In 2013, only 42 percent of German automobiles produced in US plants were actually sold in the US. The lion's share was exported from the US to either Europe or Asia. Against the backdrop of international automotive CEOs being more likely than their fellow CEOs to be "extremely concerned" about currency risks (40 percent vs. 22 percent),[22] it is straightforward to interpret currency risk or "natural" hedging as being at the heart of offshoring activities in the US. But do German automobile and mechanical engineering sectors play an extraordinary role compared to the corresponding French and UK industries in this regard? The following Table 1 tries to give an answer. It shows the current size (i.e. the number of employees rounded to hundreds) of offshore production plants in the US by the respective top-3 companies in terms of total revenues in the German, French and UK automobile and mechanical engineering sectors, respectively. Our focus is on firms listed in the respective national benchmark stock index (DAX, CAC 40, FTSE 100) that can be ascribed to one of the two key sectors.

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[22]Implying a higher awareness of the necessity to relocate plants to key currency countries in order to lower the exchange rate risk according to a recent survey among 1,201 business leaders in 69 countries conducted in the fourth quarter of 2010; see PwC (2011).

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German offshores			Size	Primary location
Mechanical engineering	(1)	*Siemens Industry*	46,000	Orlando, FL
	(2)	*Linde Group*	5,000	Murray Hill, NJ
	(3)	*ThyssenKrupp IS*	14,800	Calvert, AL
Automobile	(1)	*Volkswagen*	2,000	Chattanooga, TN
	(2)	*Daimler*	5,100	Tuscaloosa, AL
	(3)	*BMW*	8,000	Spartanburg, NC
Total			80,900	

French offshores			Size	Primary location
Mechanical engineering	(1)	*Schneider Electric*	900	Andover, MA
	(2)	*Thales*	3,000	Arlington, VA
	(3)	*Safran*	7,000	Arlington, VA
Automobile	(1)	*PSA Peugot Citroën*	–	–
	(2)	*Renault*	–	–
	(3)	*Michelin*	11,600	Greenville, SC
Total			22,500	

UK offshores			Size	Primary location
Mechanical engineering	(1)	*BAE Systems*	19,800	Arlington, VA
	(2)	*Rolls Royce Group plc*	7,000	Crosspointe, VA
	(3)	*Amec Foster Wheeler*	–	Houston, TX
Automobile	(1)	*Delphi Automotive*	5,000	Troy, MI
	(2)	*GKN Driveline*	2,000	Newton, NC
	(3)	*TI Automotive*	300	Auburn Hills, MI
Total			34,100	

Table 1: Contemporary offshore production plants in the US; Source: Statista

There are several points to note with regard to the construction of Table 1 before interpreting the figures it is displaying. First, companies usually run more than just one US offshore plant. They frequently include plants that are more or less strongly engaged in joint ventures with local original equipment manufacturers (OEMs) or suppliers. For example, the French aerospace-defence-security group *Safran* runs a joint venture in equal shares with OEM *General Electric* at a plant in Rochester (NH). The 50 percent employees of it are counted in the shown figure, while we refer to Arlington (VA) as primary location, where *Safran USA* is actually headquartered.[23] It is also worth noting

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[23]Similarly, the UK *Rolls Royce* group's figure subsumes employees from a major defence plant run at Indianapolis (IN).

that the *ThyssenKrupp Industrial Solutions* employees figure includes a steel plant with 2,700 employees in Calvert (AL) that was sold in 2013. The figure for French company *Schneider Electric* misses data from a recently, i.e. in 2015, set up plant in Livermore (CA). Finally, note that both *Amec Foster Wheeler* and *TI Automotive* do not run production sites, but offshore consulting services and administration, in the US, respectively. It is also noteworthy that British vehicle producer *Aston Martin* –with total revenues amounting to GBP 470 million not top-3 in terms of revenues and, thus, not included– recently announced to plan a US offshore plant in Alabama.

Even if we consider the recent selling and setting-up of single offshore plants in the US by our considered European economies and keep in mind the snapshot character and rough classification of sectors in Table 1, the message it is conveying is clear-cut. German key sectors' US offshoring activity (about 75 to 80 thousand employees) is higher than the one by its French (about 20 to 30 thousand employees) and UK (about 35 thousand employees) couterparts taken together. This fact is suggestive for natural hedging, besides balance sheets and value added chain effects, playing a crucial role with regard to hedging currency risks for European key sectors. It can be interpreted as an idiosyncratic German way in dealing with and immunizing against exchange rate fluctuations that can, at least in parts, explain our findings. Furthermore, the awareness of the industry in general (PwC 2011) as well as the recent plans of *Aston Martin*, in particular, support this reasoning. In sum, natural hedging in the form of offshoring to key currency country locations seems to be a promising strategy to hedge the currency risk for European core industries.

7 Conclusion

Our study was motivated by presuming that looking exclusively at the reaction of exports to real exchange rate changes is not satisfactory when addressing the question whether an appreciation of the Euro causes automobile and mechanical engineering industries in the EU "pain." Estimating a trivariate VAR, we find a depreciation of the Euro vis-à-vis the US dollar to have a positive impact on exports for German and French data. The effect is, however, insignificant in the case of the confidence indicators for German series throughout and for French series in half of the specifications. Implied impulse responses show that shocks in the exchange rate lead to a significant reaction of exports, while

there might be neither a reaction of the business climate nor of order book assessments or business confidence in general. Additionally, FEVD analysis revealed that the exchange rate contributes little (in the case of France) to nearly nothing (in the case of Germany) to the dynamics of confidence indicators. If we proceed analogously with the GBP/USD exchange rate and the UK sectoral series, the picture is different as we find throughout all our specifications a significant responsiveness to exchange rate shocks. And although this responsiveness is growing over time, it is in its direction reversed to predictions of traditional and standard open-economy models (Dornbusch 1976, Svenson 2000) as well as of the "New open-economy macroeconomics" class of models (Lane 2001). The reversion, i.e. a positive impact on expected sector-specific demand triggered by a real exchange rate deprecation, can be rationalized on two grounds: (i) by an adverse balance sheet effect in case of firms' liabilities being substantially dollarized or, more generally, denominated in foreign currency paralleled by firm revenues in domestic currency (Aghion *et al.* 2000, Bahmani-Oskooee and Miteza 2003) or (ii) by the dependence on relatively more expensive imported intermediate goods that need to be acquired in foreign currency (Greenaway *et al.* 2010). Both effects, (i) and/or (ii), can outweigh the positive effect depreciation has on net sectoral exports and, consequently, reverse demand expectations.

With regard to the hypotheses set up in the introductory sections of this paper, it is only the Williamson-hypothesis [*H2*] of an an extraordinary role being played by the German key sectors that our analysis clearly confirms. In contrast, the hypothesis of an overall high exchange rate susceptibility of the two core sectors, automobiles and mechanical engineering, as claimed by lobbies and held as popular belief [*H1*], can be rejected. From a political economy perspective this raises the question why, in particular, these industries heavily lobby for low exchange rates. A straightforward answer is that immunization strategies such as natural hedging, that is, shifting production (temporarily) abroad, as in particular observed at a relatively large scale in the German automobile and related industries are costly. Finally, we find no clear-cut evidence for the Frankel-Rose-Melitz-hypothesis [*H3*] according to which the coming into being of the Eurozone should imply a joint decline in the responsiveness of members' sectoral business confidence to real exchange rate shocks. For though the German sectoral index reaction clearly appears to markedly decrease after the Euro cash changeover, we see no indication for a similar tendency in the case of French time series.

In the light of the recent debate triggered by the efforts of the Bank of Japan to strategically lower real exchange rates, our results suggest that "competitive deprecia-

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tion" is neither an option nor a threat for Germany and its core industries in the short and medium run. The same applies to the other European economies analyzed in this study showing reversed standard model reactions of expected demand.

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A   Appendix

A.1   Time series and descriptive statistics
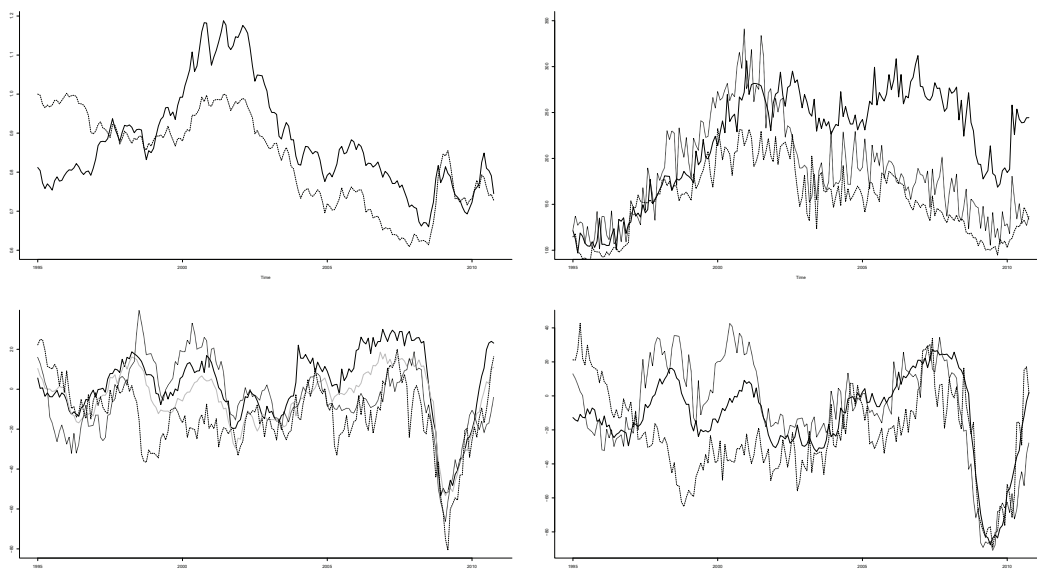
Figure 1: Clockwise from upper left corner: real EUR/USD exchange rate (solid) and real GBP/USD exchange rate (dashed), vehicle and machinery exports to U.S. for Germany (solid), UK (dashed) and France (dotted), Assessment of Export Order-Book Indicator in automobile and mechanical engineering industry for Germany (solid), UK (dashed) and France (dotted), *ifo* Business Climate (solid) as well as European Commission Confidence Indicator in automobile and mechanical engineering industry for Germany (grey), UK (dashed) and France (dotted)

Variable	Germany				France				UK			
	Min	Max	Mean	Sd	Min	Max	Mean	Sd	Min	Max	Mean	Sd
EXR	0.66	1.19	0.88	0.131	0.66	1.19	0.88	0.13	-	-	-	-
EXR_GB	-	-	-	-	-	-	-	-	0.61	1.00	0.83	0.12
EXP	96.97	312.02	219.80	57.25	101.16	341.04	185.59	53.37	85.45	232.54	154.99	37.72
BC	-53.35	30.00	1.99	17.72	-	-	-	-	-	-	-	-
COF	-53.30	18.60	-4.89	14.90	-66.30	39.65	-4.79	17.88	-80.70	24.95	-13.62	17.01
ASS_EXP	-87.60	27.25	-12.02	24.20	-90.75	42.60	-7.61	29.83	-89.15	42.65	-23.59	26.11

Table 2: Descriptive statistics for real EUR/USD exchange rate (EXR), real GBP/USD exchange rate (EXR_GB), value index of exports (EXP), ifo Business Climate (BC), European Commission Confidence Indicator (COF) and European Commission's Assessment of Export Order-Book Indicator (ASS_EXP)

Note: Prior to 2004, the business climate in the two sectors is published for West Ger-

I

many and East Germany separately. For West Germany the series of the mechanical engineering industry is subsumed under the sub-category "Investment goods without cars" (manufacturing sector), while the business climate of the automobile sector is classified under "Commodities" (manufacturing sector). For East Germany, we draw the business climate for the respective industries from the "Investment goods manufacturing industry" category. The business climate for Germany, taking East and West together, is calculated by aggregating and taking means.

A.2   Reduced form analysis of correlation structure of series

A visual inspection of the correlation function $\mathbf{P_y}(\tau)$ both for HP(129,600) and log first difference filtered series (not shown here) indicates a significant positive correlation between the series EXR and EXP as well as EXP and BC.

There is only little significant correlation between EXR and BC while there is evidence for significant cross-series dynamics between the exchange rate and exports on the one hand and exports and business climate on the other. This result is backed up and broadened by inspection of the cross-series dynamics in the frequency domain.[24] We calculate several multivariate measures in the frequency domain. If one considers two stationary time series x_t and y_t the cross-spectrum between these two is given by

$$f_{xy}(\omega) = c_{xy}(\omega) - iq_{xy}(\omega), \tag{17}$$

with $\omega \in [-\pi, \pi]$. The cospectrum $c_{xy}(\omega)$ measures the "in-phase" covariance between the two series, whereas the quadrature spectrum $q_{xy}(\omega)$ measures the covariance between the "out-of-phase" components. Together with the series autospectra this can be used to calculate the squared coherency

$$sc(\omega) = \frac{|f_{xy}(\omega)|^2}{f_x(\omega)f_y(\omega)}, \tag{18}$$

which is defined $0 \leq sc(\omega) \leq 1$. The squared coherency can be interpreted as a measure for the strength of the linear relationship between the two series at different frequencies.

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[24]The following common link exists between the time and frequency domain: $f_x(\omega) = (1/2\pi)\sum_{\tau=-\infty}^{\infty}\gamma_x(\tau)e^{-i\omega\tau}$ with $\omega \in [-\pi, \pi]$, i.e., the spectrum of a series is defined as the Fourier transform of the autocovariance function.

Unfortunately, the squared coherency does not contain information about a potential phase-shift, i.e. the lead-lag structure of the two series at different frequencies. This is why it should be interpreted jointly with the phase spectrum

$$\phi_{xy} = -arctan(q_{xy}(\omega)/c_{xy}(\omega)). \tag{19}$$

Alternatively, one may also judge the linear relationship between the two series by looking at the dynamic correlation

$$\rho(\omega) = \frac{c_{xy}(\omega)}{\sqrt{f_x(\omega)f_y(\omega)}}, \tag{20}$$

with $-1 \leq \rho(\omega) \leq 1$. The dynamic correlation measures the correlation between the "in-phase" components of x_t and y_t at frequency ω (Croux *et al.* 2001). Figure 2 shows the autospectra, integrated spectra, amount of explained variance and dynamic correlation computed from bivariate VARs fitted to our vector of observations $\boldsymbol{y}_t$ for the HP and logD filtered series, respectively.
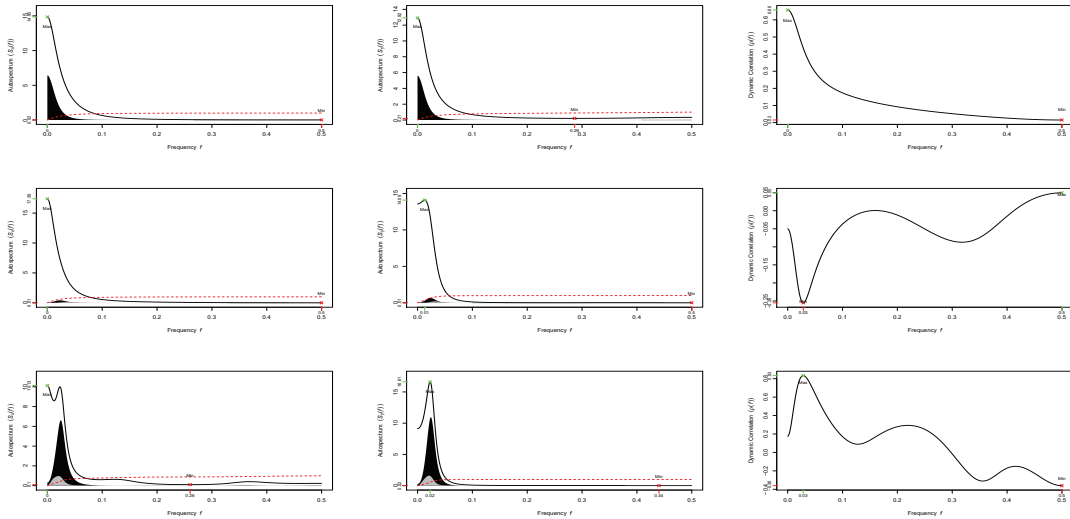
Figure 2. Bivariate spectral properties: EXR, EXP, BC; filter: HP

Note: First row – EXR vs. EXP, second row – EXR vs. BC,
third row – EXP vs. BC

The first row of Figure 2 shows the autospectra of the real EUR/USD exchange rate and exports as well as the dynamic correlation between the series. Looking at the series' autospectra, we see that quite a significant part of the exports' variance is explained by

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exchange rate dynamics and vice versa.[25] Furthermore, the two series are highly and positively correlated at the most relevant frequencies ($\rho_{max} = 0.66$).[26]

Things look somewhat different in case of the intra-series dynamics between the exchange rate and the business climate (second row of graphs). It can be seen that only a negligible part of the business climate's variance can be explained by the EUR/USD exchange rate. Additionally, the absolute value of the dynamic correlation at the relevant frequencies is much smaller. Looking at the last row of Figure 2 we see that there is a strong relationship between exports and the business climate in terms of explained variance and profound dynamic correlation. The latter takes on a maximum value of $\rho = 0.83$ around the dominating low frequencies. Though less pronounced, the results for using the logD-filter are, in general, qualitatively in line with this picture. Detailed results are available on request from the authors.

A.3 Gibbs sampler

This section briefly sketches the Gibbs sampling algorithm. A more detailed explanation of the different sampling steps can be found in Primiceri (2005).

A.3.1 Step 1: Initialization

Initialize $\boldsymbol{A}^T$, $\boldsymbol{B}^T$, $\boldsymbol{\Sigma}^T$ and the hyperparameters Q, W, S.

A.3.2 Step 2: Drawing coefficient states $\boldsymbol{B}^T$

Conditional on the data and all other parameters the observation equation 6 is linear and has Gaussian innovations. $\boldsymbol{B}^T$ is sampled from $p(\boldsymbol{B}^T|\boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, W, S)$. Based on the algorithm of Carter and Kohn (1994) draws for $\boldsymbol{B}_t = \boldsymbol{B}_{t-1} + \nu_t$ are obtained from $p(\boldsymbol{B}^T|\boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, W, S)$, which is $N(\boldsymbol{B}_{t|t+1}, \boldsymbol{P}_{t|t+1})$ with
$\boldsymbol{B}_{t|t+1} = E(\boldsymbol{B}_t|\boldsymbol{B}_{t+1}, \boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, W, S)$,
$\boldsymbol{P}_{t|t+1} = Var(\boldsymbol{B}_t|\boldsymbol{B}_{t+1}, \boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, W, S)$.

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[25]The explained variance is given by the colored area (black shaded area – "in-phase" component, grey shaded area – "out-of-phase" component of explained variance).

[26]"Relevant" in the sense of these frequencies covering most of the series' variance.

$\boldsymbol{B}_{t|t+1}$ and $\boldsymbol{P}_{t|t+1}$ are calculated using the forward filtering, backward sampling (FFBS) algorithm.

A.3.3 Step 3: Drawing covariance states $\boldsymbol{A}^T$

Together equation 6 and $\boldsymbol{X}'_t = \boldsymbol{I}_n \otimes [1, \boldsymbol{y}'_{t-1}, \boldsymbol{y}'_{t-2}, ..., \boldsymbol{y}'_{t-p}]$ can be rewritten as $\boldsymbol{A}_t(\boldsymbol{y}_t - \boldsymbol{X}'_t\boldsymbol{B}_t) = \boldsymbol{A}_t\hat{\boldsymbol{y}}_t = \boldsymbol{\Sigma}_t\boldsymbol{\epsilon}_t$, where, given $\boldsymbol{B}^T$, $\hat{\boldsymbol{y}}_t$ is observable. Under the assumption that S is block-diagonal the Carter and Kohn (1994) algorithm can be used again to draw $\alpha_{i,t}$ from $N(\alpha_{i,t|t+1}, \Lambda_{i,t|t+1})$, where

$\alpha_{i,t|t+1} = E(\alpha_{i,t}|\alpha_{i,t|t+1}, \boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, W, S)$ and

$\Lambda_{i,t|t+1} = Var(\alpha_{i,t}|\alpha_{i,t|t+1}, \boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{\Sigma}^T, Q, W, S)$.

See Primiceri (2005) for a discussion of the problem of non-linearity and the assumptions made with regard to S in drawing $\boldsymbol{A}^T$.

A.3.4 Step 4: Drawing volatility states $\boldsymbol{\Sigma}^T$

In a next step we consider $\boldsymbol{A}_t(\boldsymbol{y}_t - \boldsymbol{X}'_t\boldsymbol{B}_t) = \boldsymbol{y}^*_t = \boldsymbol{\Sigma}_t\boldsymbol{\epsilon}_t$. Now, given $\boldsymbol{A}^T$ and $\boldsymbol{B}^T$, $\boldsymbol{y}^*_t$ can be observed. Squaring and taking logs transforms this system of equations into a linear one: $\boldsymbol{y}^{**} = 2\boldsymbol{h}_t + \boldsymbol{e}_t$ with $\boldsymbol{h}_t = \boldsymbol{h}_{t-1} + \boldsymbol{\eta}_t$, where $y^{**}_{i,t} = log((y^*_{i,t})^2 + \bar{c})$, $\bar{c} = 0.001$ is an offset constant that makes the estimation procedure more robust, $e_{i,t} = log(\epsilon^2_{i,t})$, $h_{i,t} = log(\sigma_{i,t})$. Although, the system is now linear, its innovations are non-Gaussian since the errors are $\sim log\chi^2(1)$. Following Kim *et al.* (1998) we use a mixture of seven normal densities for each element of $\boldsymbol{e}$. The seven normal densities have component probabilities q_j, means $m_j - 1.2704$ and variances v^2_j, with $j = 1, ..., 7$. If we define $s^T = [s_1, ..., s_T]'$ to be the matrix of indicator variables that select for each element of $\boldsymbol{e}_t$ the respective member of the mixture, conditional on $\boldsymbol{B}^T$, $\boldsymbol{A}^T$, Q, W, S and s^T the system is now linear and approximately Gaussian. Thus, the Carter and Kohn (1994) algorithm can be used again to recursively draw the volatilities $\boldsymbol{h}_t$ from $N(\boldsymbol{h}_{t|t+1}, \boldsymbol{H}_{t|t+1})$, with $\boldsymbol{h}_{t|t+1} = E(\boldsymbol{h}_t|\boldsymbol{h}_{t+1}, \boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{B}^T, Q, W, S, s^T)$ and $\boldsymbol{H}_{t|t+1} = Var(\boldsymbol{h}_t|\boldsymbol{h}_{t+1}, \boldsymbol{y}^T, \boldsymbol{A}^T, \boldsymbol{B}^T, Q, W, S, s^T)$.

j	q_j	m_j	v_j^2
1	0.0073	-10.12999	5.79596
2	0.10556	-3.97281	2.61369
3	0.00002	-8.56686	5.17950
4	0.04395	2.77786	0.16735
5	0.34001	0.61942	0.64009
6	0.24566	1.79518	0.34023
7	0.25750	-1.08819	1.26261

Table 3: Mixing Distributions as in Kim *et al.* (1998)

A.3.5 Step 5: Drawing hyperparameters Q, W and S

Conditional on $\boldsymbol{y}^T$, $\boldsymbol{B}^T$, $\boldsymbol{A}^T$ and $\boldsymbol{\Sigma}^T$ the hyperparameters Q, W and S can be obtained directly from their inverse-Wishart posterior distributions.

A.3.6 Step 6: Sampling replications

Go back to Step 2.

A.4 Impulse responses and FEVD
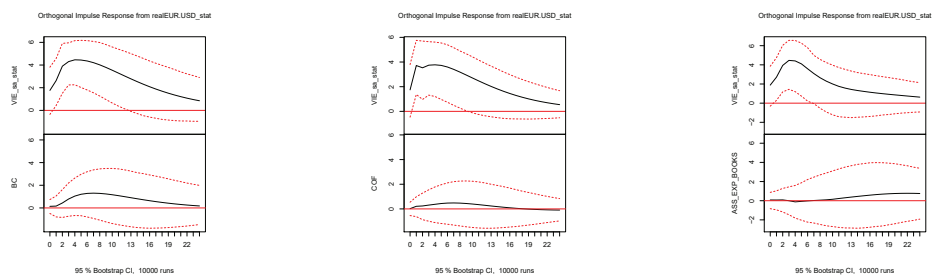
Figure 3: Impulse responses of EXP and respective climate indicator for orthogonal shock in EXR (HP-filtered series)
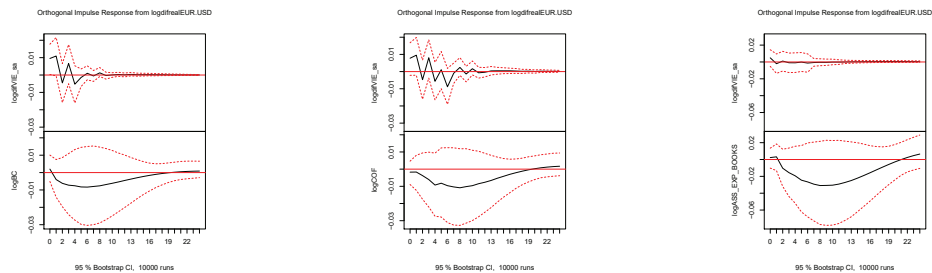
Figure 4: Impulse responses of EXP and respective climate indicator for orthogonal shock in EXR (logD-filtered series)

Note:

The underlying VAR models of the first column of figures on this page use the ifo business climate indicator, of the second column the EC confidence indicator, and of the third column the EC assessment of export order-book indicator as third variable in endogenous vector, respectively. All indicators refer to the automobile and mechanical engineering industries only.
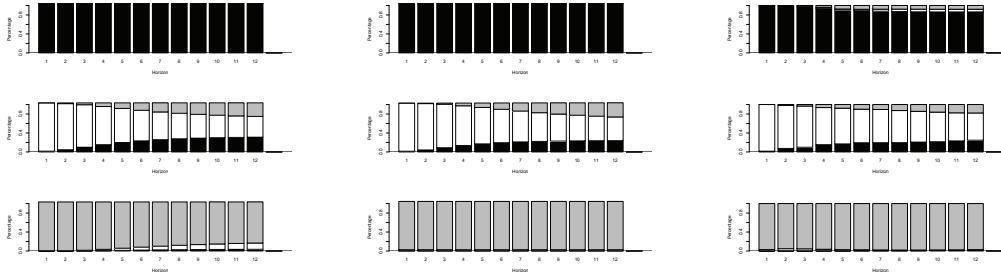
Figure 5: FEVD for EXR, EXP and respective climate indicator (HP)
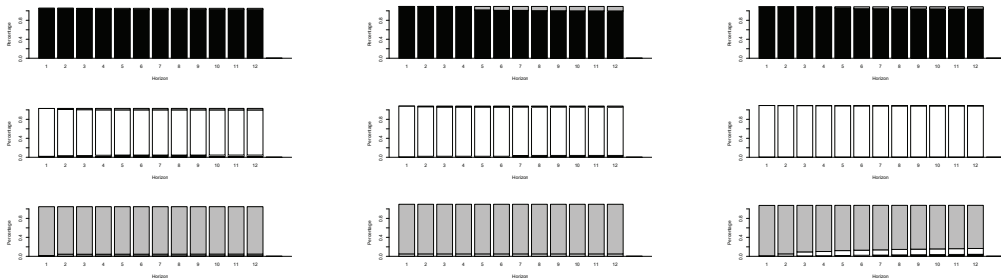
Figure 6: FEVD for EXR, EXP and respective BC indicator (logD)

Note:

The underlying VAR models of the first column of figures on this page use the ifo business climate indicator, of the second column the EC confidence indicator, and of the third column the EC assessment of export order-book indicator as third variable in endogenous vector, respectively. All indicators refer to the automobile and mechanical engineering industries only.
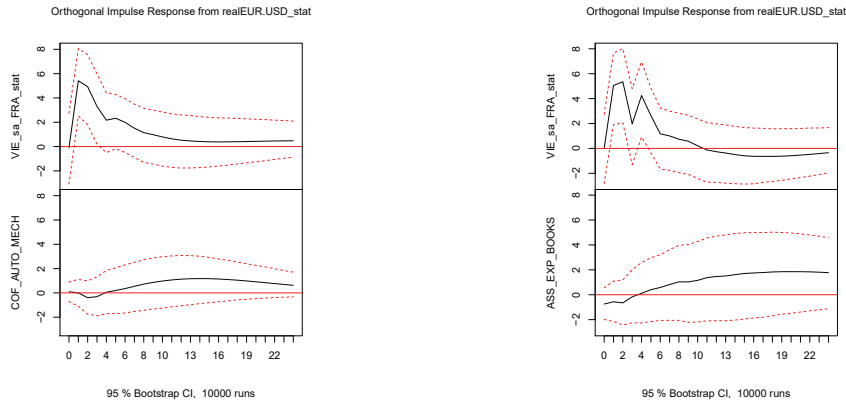
Figure 7: Impulse responses of EXP and respective climate indicator for orthogonal shock in EXR (HP): France
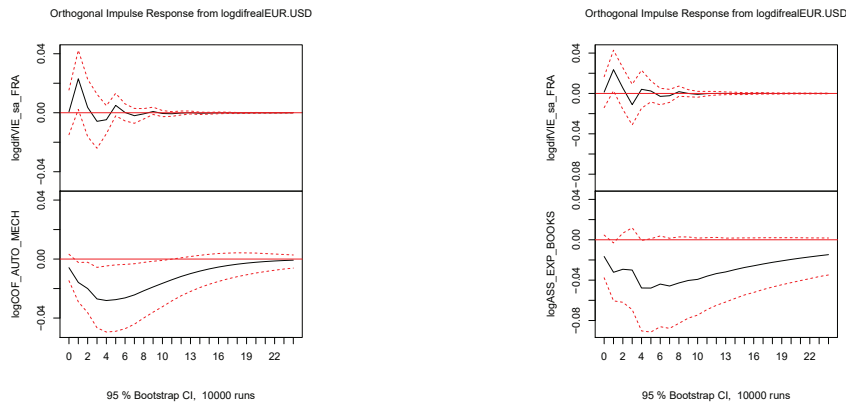
Figure 8: Impulse responses of EXP and respective BC indicator for orthogonal shock in EXR (logD): France

Note:

The underlying VAR models of the first column of figures on this page use the EC confidence indicator, the second column the EC assessment of export order-book indicator, respectively. All indicators refer to the automobile and mechanical engineering industries only.
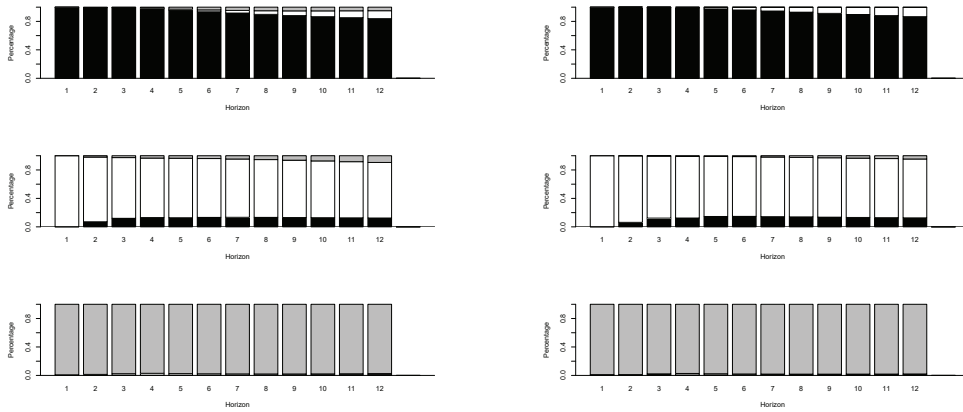
Figure 9: FEVD for EXR, EXP and respective climate indicator (HP): France
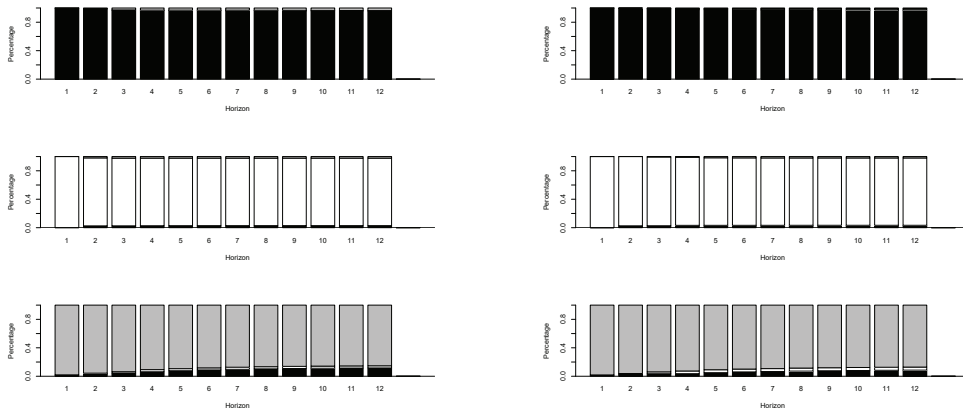
Figure 10: FEVD for EXR, EXP and respective climate indicator (logD): France

Note:

The underlying VAR models of the first column of figures on this page use the EC confidence indicator, the second column the EC assessment of export order-book indicator, respectively. All indicators refer to the automobile and mechanical engineering industries only.
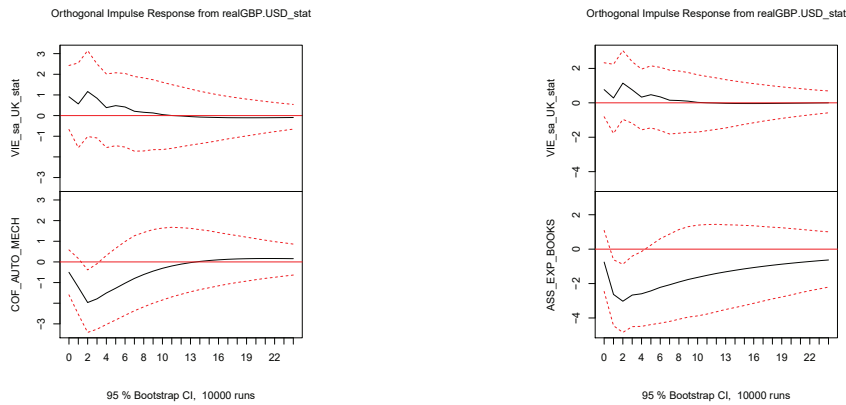
Figure 11: Impulse responses of EXP and respective climate indicator for orthogonal shock in EXR (HP): UK
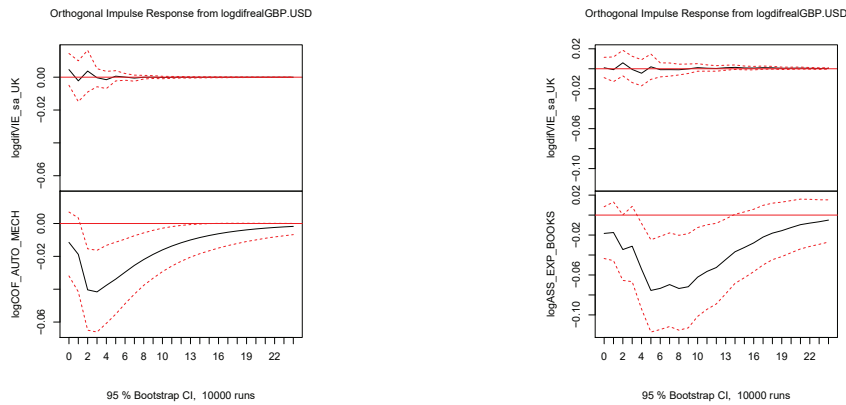
Figure 12: Impulse responses of EXP and respective climate indicator for orthogonal shock in EXR (logD): UK

Note:

The underlying VAR models of the first column of figures on this page use the EC confidence indicator, the second column the EC assessment of export order-book indicator, respectively. All indicators refer to the automobile and mechanical engineering industries only.
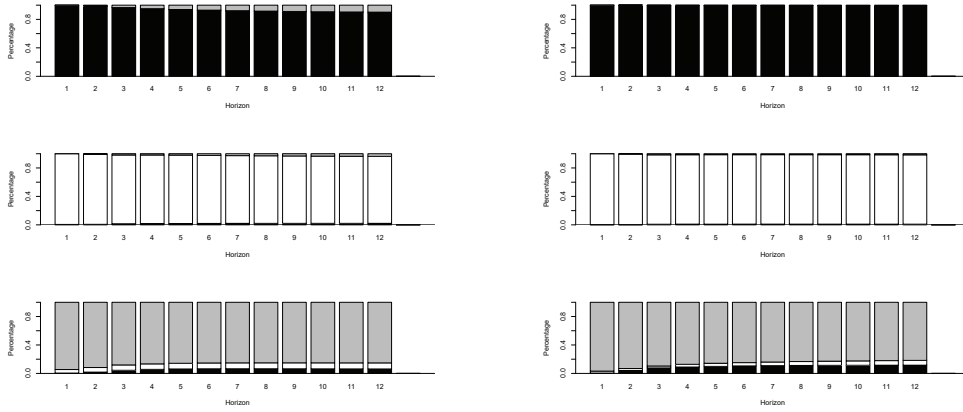
Figure 13: FEVD for EXR, EXP and respective climate indicator (HP): UK
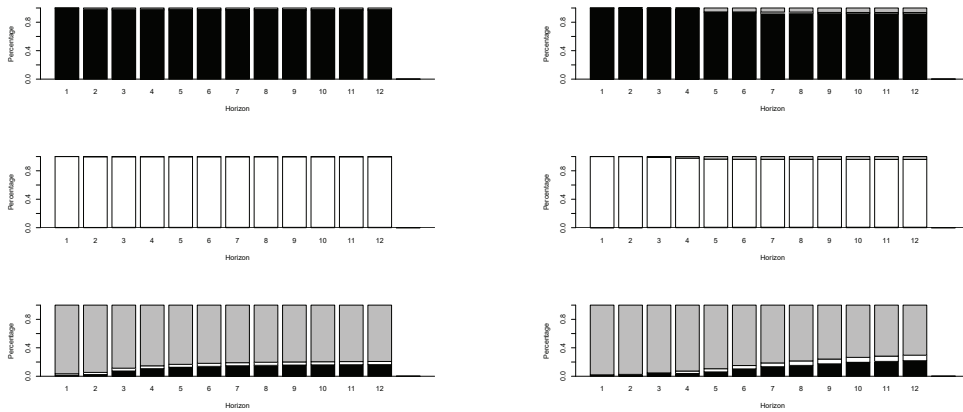
Figure 14: FEVD for EXR, EXP and respective climate indicator (logD): UK

Note:

The underlying VAR models of the first column of figures on this page use the EC confidence indicator, the second column the EC assessment of export order-book indicator, respectively. All indicators refer to the automobile and mechanical engineering industries only.

A.5 Estimation results

Endogenous variables: EXR, EXP, BC
Deterministic variables: const
Method: Ordinary Least Squares
Sample: Jan1995 Oct2010
Sample size: 187
Log Likelihood: -806.912

Equation: $EXR = EXR.l1 + EXP.l1 + BC.l1 + EXR.l2 + EXP.l2 + BC.l2 + EXR.l3 + EXP.l3 + BC.l3 + const$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
EXR.l1	$1.287e+00$	$7.494e-02$	17.179	$< 2e-16$***
EXP.l1	$1.438e-04$	$1.067e-04$	1.347	0.1796
BC.l1	$-3.613e-05$	$3.380e-04$	-0.107	0.9150
EXR.l2	$-5.535e-01$	$1.149e-01$	-4.815	$3.14e-06$***
EXP.l2	$1.445e-04$	$1.091e-04$	1.324	0.1871
BC.l2	$-6.117e-04$	$4.868e-04$	$4.868e-04$	0.2106
EXR.l3	$1.569e-01$	$7.582e-02$	2.070	0.0399*
EXP.l3	$-1.234e-04$	$1.084e-04$	-1.138	0.2566
BC.l3	$4.566e-04$	$3.392e-04$	1.346	0.1799
const	$3.582e-04$	$1.464e-03$	0.245	0.8070
Multiple R-squared	0.8975	Adj. R-squared		0.8923
F-statistic	172.2 on 9 and 177 DF	p-value		<2.2e-16

Table 4: Estimation Results VAR[3]-Model

Equation: $EXP = EXR.l1 + EXP.l1 + BC.l1 + EXR.l2 + EXP.l2 + BC.l2 + EXR.l3$
$+ EXP.l3 + BC.l3 + const$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
EXR.l1	$1.637e + 02$	$5.314e + 01$	3.080	0.002399**
EXP.l1	$2.858e - 01$	$7.568e - 02$	3.777	0.000217***
BC.l1	$4.932e - 01$	$2.397e - 01$	2.058	0.041070*
EXR.l2	$-1.018e + 02$	$8.151e + 01$	-1.249	0.213151
EXP.l2	$2.234e - 01$	$7.737e - 02$	2.887	0.004374**
BC.l2	$-3.598e - 02$	$3.452e - 01$	-0.104	0.917115
EXR.l3	$3.624e + 01$	$5.377e + 01$	0.674	0.501174
EXP.l3	$7.433e - 03$	$7.688e - 02$	0.097	0.923089
BC.l3	$-5.433e - 02$	$2.405e - 01$	-0.226	0.821541
const	$-7.250e - 01$	$1.038e + 00$	-0.698	0.485860
Multiple R-squared	0.6426		Adj. R-squared	0.6244
F-statistic	35.36 on 9 and 177 DF		p-value	<2.2e-16

Equation: $BC = EXR.l1 + EXP.l1 + BC.l1 + EXR.l2 + EXP.l2 + BC.l2 + EXR.l3$
$+ EXP.l3 + BC.l3 + const$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
EXR.l1	-12.78167	16.01834	-0.798	0.42597
EXP.l1	0.03538	0.02281	1.551	0.12264
BC.l1	1.02815	0.07224	14.233	$< 2e - 16$***
EXR.l2	41.99795	24.56937	1.709	0.08914.
EXP.l2	-0.06798	0.02332	-2.915	0.00402**
BC.l2	0.20088	0.10405	1.931	0.05511.
EXR.l3	-12.96744	16.20730	-0.800	0.42473
EXP.l3	-0.03612	0.02317	-1.559	0.12086
BC.l3	-0.22625	0.07250	-3.121	0.00211**
const	0.06351	0.31291	0.203	0.83938
Multiple R-squared	0.947		Adj. R-squared	0.9443
F-statistic	351.3 on 9 and 177 DF		p-value	<2.2e-16

Table 5: Estimation Results VAR[3]-Model cont'ed

Endogenous variables: EXR, EXP, BC
Deterministic variables: const
Method: Ordinary Least Squares
Sample: Jan1995 Oct2010
Sample size: 187
Log Likelihood: -817.391

Equation: $EXR = EXR.l1 + EXR.l2 + EXR.l3$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
EXR.l1	1.31033	0.07320	17.901	$< 2e-16$***
EXR.l2	-0.53173	0.11495	-4.626	$7.02e-06$***
EXR.l3	0.14730	0.07324	2.011	0.0457*
Multiple R-Squared	0.8924	Adjusted R-squared		0.8907
F-statistic	508.8 on 3 and 184 DF		p-value	<2.2e-16

Equation: $EXP = EXR.l1 + EXP.l1 + BC.l1 + EXP.l2$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
EXR.l1	102.54018	22.34386	4.589	$8.24e-06$***
EXP.l1	0.28481	0.07148	3.985	$9.75e-05$***
BC.l1	0.41549	0.07809	5.321	$3.00e-07$***
EXP.l2	0.21449	0.06836	3.138	0.00198**
Multiple R-Squared	0.6376	Adjusted R-squared		0.6297
F-statistic	80.5 on 4 and 183 DF		p-value	<2.2e-16

Equation: $BC = BC.l1 + EXR.l2 + EXP.l2 + BC.l3$

Variable	Coefficient	Std. Error	t-Statistic	Prob.
BC.l1	1.15062	0.04566	25.198	$< 2e-16$***
EXR.l2	18.87177	6.74269	2.799	0.005678**
EXP.l2	-0.07381	0.02065	-3.575	0.000448***
BC.l3	-0.13906	0.05005	-2.778	0.006036**
Multiple R-Squared	0.945	Adjusted R-squared		0.9438
F-statistic	786.6 on 4 and 183 DF		p-value	<2.2e-16

Table 6: Estimation Results Subset VAR[3]-Model

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