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A Disaggregated Perspective

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Joscha Beckmann, Ansgar Belke, and Florian Verheyen[1]

Exchange Rate Pass-through into German Import Prices – A Disaggregated Perspective

Abstract

This study analyzes the exchange rate pass-through into German import prices based on disaggregated data taken on a monthly basis between 1995 and 2012. Our main contribution is twofold: firstly, we employ various time-series techniques to analyze data for different product categories, and also cointegration techniques to carefully distinguish between shortrun and long-run pass-through coefficients. Secondly, in a panel data approach we estimate time-varying pass-through coefficients and explain their development with regard to various macroeconomic factors. Our results show that long-run pass-through is only partly observable and incomplete, while short-run pass-through shows a more unique character, although heterogeneity across product groups does exist. We are also able to identify several macroeconomic factors which determine changes in the degree of pass-through, which is especially relevant for policymakers.

JEL Classification: E31, F10, F14

Keywords: Exchange rate pass-through; germany; cointegration; time-varying coefficient model

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1. Introduction

Although the concept of exchange rate pass-through (ERPT) is straightforward from a theoretical point of view, economic research has struggled remarkably when it comes to explaining the ambiguous empirical results yielded by different studies. In a nutshell, the elasticities of import prices with regard to changes in the exchange rate can range from 0% to 100%, depending on the pricing strategy of exporters (Brun-Aguere et al., 2012). More precisely, everything depends on whether foreign exporters rely on a strategy of producer-currency pricing (PCP) or local-currency pricing (LCP) (Campa and Goldberg, 2005). The former case corresponds to the law of one price, which is based on the assumption that currency fluctuations directly translate into import prices. By contrast, the latter denotes a situation where producers apply different pricing strategies in different markets. Overall, the empirical evidence suggests that pass-through is incomplete. This finding can be attributed to a price-setting mechanism based on time-varying mark-ups over marginal costs and other macro- and microeconomic factors such as monetary policy or market concentration (Brun-Aguere et al., 2012). Such price mark-ups can influence the degree of ERPT when markets deviate from the perfect competition framework. In a situation with imperfect competition, producers will set prices using a mark-up over their marginal costs (Menon, 1996). These mark-ups can of course be adjusted to gain market shares, which would influence the degree of pass-through.

However, while several studies have analyzed the degree of ERPT from various aspects, there is no consensus on which factors may be responsible for changes in the degree of pass-through. Generally speaking, the degree of pass-through evolves over time and also differs over time horizons. Factors which are discussed to explain this pattern include shifts in the composition of imported goods and changing competition (Stahn, 2009).[3] In this vein, different pass-through elasticities across industries may also provide an explanation for a changing pass-through into aggregate import prices over time (Campa and Goldberg, 2010). Finally, an interesting question is whether importers and exporters follow an identical price-setting strategy, since the corresponding elasticities are only identical in a symmetrical two-country model (Stahn, 2009). The case of Germany is of particular importance in this context: on the one hand it is among the leading exporters, on the other, changes in German import prices may be transmitted not only into German consumer prices but also into the price level of the eurozone as a whole, owing to Germany's leading position.

[3] See Stahn (2009) for a discussion of the question of whether this wisdom also holds with regard to disaggregated data.

In this context, the present study focuses on pass-through in relation to German import prices. Overall, results suggest that the sensitivity of ERPT significantly differs between the short- and the long-run and decreases over time (Stahn, 2009). Against this background, we analyze the exchange rate pass-through for import price data between 1995 and 2012.[4] By analyzing both the individual product categories in the time-series dimension and the general character of pass-through in a panel framework our main contribution is twofold: firstly, we analyze data for different product categories (thus avoiding the so-called "aggregation bias") using cointegration techniques and carefully distinguish between short-run and long-run pass-through coefficients. Secondly, we use panel techniques to evaluate overall changes in the pass-through mechanism. We estimate time-varying pass-through coefficients and try to explain their development according to different macroeconomic factors. Overall, this enables us to analyze the difference in pass-through across product categories and time horizons while also tackling the question of whether the evolution of pass-through coefficients is driven by common macroeconomic factors across product groups.

From the point of view of monetary policy, a closer look at the degree and character of pass-through is necessary to evaluate a possible inflation threat. In addition, important implications arise for domestic producers in terms of price calculation and production structure. The fact that monetary policymakers are interested in a more general analysis of the pass-through mechanism while producers might pay more attention to specific product groups demonstrates the usefulness of analyzing both individual product categories and common factors in terms of pass-through. Evaluating both perspectives is also important since incomplete pass-through has been identified as a reason for sloppy exchange rate adjustment in the context of global current account imbalances. (Obstfeld and Rogoff, 2005).

The remainder of this paper is organized as follows: the following section summarizes previous empirical findings related to our study. Section 3 begins with a description of our data and the introduction of our modeling cycle. As a first step in our empirical analysis, we consider both a multivariate and a single-equation cointegration approach, where we analyze the degree of long-run pass-through and the evolution of the coefficients. We then turn to short-run models in first differences, where the coefficients of the variables are also allowed to change over time according to two different frameworks. In a final step, we leave the time series dimension and analyze the general evolution of the pass-though coefficients in a panel

[4] See, for example, Mumtaz et al. (2006) for a disaggregated analysis of ERPT into UK import prices. They find considerable variation in the ERPT across fifty seven industries.

framework, using different macroeconomic variables as common explanatory variables. After a presentation of the empirical results, Section 4 concludes.

2. Literature review

Considering the enormous amount of research that has focused on exchange rate pass-through, we have decided to direct our attention to previous research that is closely related to our study. Roughly three different strands in the literature can be distinguished. While early studies focused on the question of whether pass-through is complete or incomplete, recent research has analyzed two different questions: firstly, whether micro- or macro-factors determine the degree of ERPT, and secondly, whether pass-through is endogenous or not (Brun-Aguere et al., 2012). However, since our approach intends to analyze the underlying dynamics in a fairly general fashion we aim at addressing all three questions simultaneously, although we do not consider micro-factors as explanatory variables in our approach because of the data limitations.

Theoretical explanations for an incomplete ERPT can be found in a seminal paper by Dornbusch (1987). In this paper, he presents a model which explains the extent of ERPT by, for example, product substitutability or the degree of competition prevailing in the market. While Dornbusch (1987) attributes an incomplete ERPT to micro-economic determinants, Taylor (2000) argues that the recent decline in pass-through is due to the decline in overall inflation. This hypothesis is supported, for example, by Bailliu and Fujii (2004) for a sample of eleven industrialized countries, or Ben Cheikh (2012a), who focuses on twenty seven OECD countries. In contrast - at least to some extent - Campa and Goldberg (2005) attribute the decline in pass-through particularly to a change in the composition of imports. As imports consist less of energy and ever more of manufactured products, pass-through declines in many countries. This change in import composition is superior to macro-economic effects. However, this view is challenged by Bussière and Peltonen (2008). They find that ERPT is especially determined by macro-economic factors, which might legitimize us considering only macro-economic determinants.

Turning to recent results for German data, Bussière (2007) estimates pass-through coefficients for G7 countries. While his results point to nonlinearities and asymmetries in the ERPT mechanism, he finds that the exchange rate elasticity of German import prices is rather low compared with the other countries in the sample.[5] To be precise, in the linear model he finds a

[5] In general, pass-through is supposed to be larger in small open economies (Ghosh, 2012).

short-run coefficient of -0.33 and a long-run coefficient of -0.36.[6] De Bandt et al. (2007) apply time-series and panel data techniques to EMU data disaggregated by SITC categories. Furthermore, they allow for structural breaks in the relationship. The long-run pass-through is found to be incomplete in seven out of nine cases for Germany. However, the estimates for Germany do not generally stand out with respect to its magnitude. Bussière and Peltonen (2008) estimate export and import price equations for forty one countries using data in first differences. Accordingly, they do not consider cointegration. For Germany, they find a short-run coefficient of 0.31. Again, the German coefficient is in the middle of the estimates rather than being located at the one or the other extreme. Stahn (2009) performs a disaggregated analysis for Germany. From her analysis of eleven product categories she finds that the pass-through has decreased. She merges data from two different product classifications, as there was a break in 1995. In the present study we will avoid potential problems of aggregating data from two different statistics, since we rely solely on data from 1995 onward. Furthermore, we rely on a broader disaggregation as compared to Stahn (2009).

Berner (2011) investigates the pass-through into German unit value indices. While he considers a very detailed dataset at the 8-digit level, his evidence points to an average pass-through of 0.42 after three months and 0.46 after nine months. However, owing to mixed evidence of stationarity, he does not consider cointegration and his data relates mostly to electronic and textile products. In contrast, our analysis will give a more complete picture, as we consider the whole spectrum of German imports. In addition, he distinguishes the pass-through across trading partners. The pass-through seems to be low, especially for imports from European countries.

In a recent study, Delatte and López-Villavicencio (2012) consider asymmetries in the ERPT mechanism for Japan, Germany, the UK and the US. Their results suggest that the pass-through to consumer prices is stronger for depreciations than for appreciations. For Germany, they find asymmetries in both the short- and the long-run. A depreciation of 10% results in a rise in German consumer prices of 1.2%. The effect of appreciation is insignificant. Brun-Aguerre et al. (2012) estimate a number of models which differ with respect to the lag length or the estimation period in a broad sample of emerging and industrialized countries. According to their results, the short-run pass-through is on average around 0.59 and the long-run effect around 0.70.

[6] One should note that the sign of the coefficient should be either positive or negative depending on whether the exchange rate is denominated in direct or indirect quoting. In general, an appreciation in the German currency should be followed by a reduction in import prices and vice versa.

Finally, two investigations by Ben Cheikh (2012a, 2012b) apply a panel cointegration approach to twenty seven OECD countries. When estimates for individual countries are considered he finds that there seems to be heterogeneity among them. For Germany the long-run pass-through is 0.63, which is significantly different from zero but not from one. Ben Cheikh (2012b) evaluates various forms of asymmetry in the ERPT mechanism in conjunction with consumer price inflation in twelve EMU countries. For example, he finds nonlinearities regarding the size and direction of a change in the exchange rate or when the economy is in an up- or downswing. For the linear model, the coefficient estimates are 0.05 (short-run) and 0.12 (long-run), which are both at the lower end of the values compared with the other eleven EMU countries.

For the methodological view, Kiliç (2010) considers smooth transition models to allow for nonlinearity and threshold effects. From an evaluation of import price equations similar to that in our study, he finds a linear short-run (long-run) coefficient of 0.57 (0.91), where the long-run coefficient is not statistically different from one. Regarding nonlinearities, the evidence in favor of nonlinear dynamics is weaker for Germany compared to other countries. For this reason and owing to the comparably short sample under investigation, we do not consider endogenously determined threshold or stochastic regime-switching nonlinearities in our empirical setup.[7] However, by applying recursive estimation techniques, we allow for nonlinearities in terms of changes in the short- and long-run pass-through mechanism. Campa and Goldberg (2010) estimate varying pass-through coefficients for twenty one industrialized countries. They find a declining pass-through across the distribution chain. While the pass-through into border prices is 0.80, that into the CPI is only 0.07.

To sum up, a number of studies investigate the ERPT into German (import) prices. While the pass-through is usually found to be incomplete both in the short- and the long-run, most of these investigations, with the exception of de Bandt et al. (2007), Stahn (2009) and Berner (2011), do not use disaggregated price indices. Generally speaking, our study differs from previous research on German import prices in several respects: in order to improve on Stahn (2009) we will use a broader disaggregation and fully comparable data. In distinction to Berner (2011) we will consider cointegration and use real price indices rather than unit value indices. De Bandt et al. (2007) also use unit value indices. Furthermore, our approach provides a deeper understanding of the forces that drive the differences and evolution of pass-

[7] Bussière (2007) also detects only weak signs of nonlinearity for German ERPT. However, Al-Abri and Goodwin (2009), or Delatte and López-Villavicencio (2012), do find evidence of nonlinearities for ERPT into German prices. Nevertheless, Delatte and López-Villavicencio (2012) note that their results for Germany might be special owing to their short sample, which is from 1991 to 2009.

through coefficients. On an international level, the work by Mumtaz et al. (2006) is probably the closest to our work, both conceptually and methodologically. However, the fact that Mumtaz et al. (2006) do not consider cointegration still represents a major difference from our study.

3. Empirical analysis

3.1 Data

The import price data is provided by the Federal Statistical Office of Germany and consists of twenty one product categories of 2-digit accuracy, according to German Product Classification (Güterverzeichnis für Produktionsstatistiken, GP). Table 1 provides a detailed description of the import price indices. The categories were chosen according to data availability. In order to have a balanced sample, we have considered only product categories for which data is available from January 1995. Our sample ends in April 2012. In total, the share of the twenty one import price subindices accounts for almost 60% of the overall German import price index. As can be seen from Table 1, there are considerable differences with regard to the average annual inflation rates and their volatility across the categories. While total import price inflation has been on average 1.81% per year, which is consistent with the definition of price stability of the ECB, raw materials such as oil and ores exhibit much stronger increases. Furthermore, the volatility of these categories (measured as the standard deviation in the annual inflation rates) is also larger. However, from a policymaker's point of view there is no indication that categories with higher inflation rates or higher volatility have larger shares in the total import price index. We take these descriptive statistics as a first piece of evidence that a disaggregated analysis is appropriate.

Table 1: Descriptive statistics

Product category	Share in ‰	Average inflation	Variability of inflation
GP00: total import price index	1000	1.81	5.17
GP02: forestry and logging	0.79	3.17	8.73
GP03: fishing and aquaculture	0.83	2.39	11.73
GP05: coal and peat	3.90	7.60	21.95
GP06: crude oil and natural gas	92.50	15.56	31.58
GP07: mineral ores	6.60	11.38	19.68
GP08: stone and other mining products	2.33	2.07	2.17
GP11: beverages	6.62	1.40	1.76
GP12: tobacco products	1.27	1.13	2.18
GP15: leather and leatherware	11.25	1.53	2.22
GP16: wood; basketry; wattle and cork works (without furniture)	7.83	0.85	4.92
GP17: paper, pasteboard and goods made thereof	22.26	0.69	6.91
GP19: cookery and mineral oil products	26.38	15.41	32.39
GP21: pharmaceutical goods	34.70	0.94	2.92
GP22: rubber and plastic goods	27.39	0.80	2.30
GP23: glass and glassware, pottery and processed stone	11.64	0.85	1.25
GP24: metals	64.43	4.94	13.69
GP25: metal products	26.83	1.33	1.88
GP28: machinery	80.78	1.13	1.20
GP29: cars and car parts	116.53	0.62	0.62
GP30: other vehicles	5.85	1.21	2.68
GP31: furniture	13.06	1.06	1.10

Nominal and real effective exchange rates calculated on unit labor costs, as well as the index of industrial production and the producer price index of Germany are all provided by the IMF. Price data is seasonally adjusted using the Census X-12 procedure. The index of industrial production is already deseasonalized. For reliable statistical inference in the following empirical analysis it is necessary to check whether the time series under consideration are stationary or not. Especially for the cointegration analysis, all time series under consideration have to be integrated of the same order, usually of order one (I(1)). From our results in Table 2 we have robust evidence that all time series are non-stationary in levels. Both augmented Dickey-Fuller and Phillips-Perron tests with various specifications do not reject the null hypothesis of a unit root when considering the time series in levels. However, transforming them into first differences yields stationary series. Therefore, a cointegration approach based on I(1) data is the most adequate framework.

Table 2: Unit root tests

Time series	No constant, no trend				Constant				Constant and trend			
	ADF test		PP test		ADF test		PP test		ADF test		PP test	
	level	diff.	level	diff.	level	diff.	level	diff.	Level	diff.	Level	diff.
GP02	0.97	0.00	0.96	0.00	0.43	0.00	0.43	0.00	0.69	0.00	0.62	0.00
GP03	0.79	0.00	0.77	0.00	0.52	0.00	0.50	0.00	0.15	0.00	0.21	0.00
GP05	0.94	0.00	0.93	0.00	0.88	0.00	0.88	0.00	0.53	0.00	0.48	0.00
GP06	0.97	0.00	0.98	0.00	0.83	0.00	0.84	0.00	0.06	0.00	0.21	0.00
GP07	0.99	0.00	0.99	0.00	0.96	0.00	0.96	0.00	0.57	0.00	0.34	0.00
GP08	1.00	0.00	1.00	0.00	1.00	0.00	1.00	0.00	0.99	0.00	0.97	0.00
GP11	1.00	0.00	1.00	0.00	0.91	0.00	0.96	0.00	0.28	0.00	0.45	0.00
GP12	0.99	0.00	0.99	0.00	0.78	0.00	0.79	0.00	0.64	0.00	0.63	0.00
GP15	1.00	0.00	1.00	0.00	0.93	0.00	0.95	0.00	0.86	0.00	0.88	0.00
GP16	0.83	0.00	0.83	0.00	0.16	0.00	0.57	0.00	0.02	0.01	0.26	0.00
GP17	0.75	0.00	0.81	0.00	0.15	0.00	0.33	0.00	0.07	0.00	0.29	0.00
GP19	0.96	0.00	0.96	0.00	0.78	0.00	0.79	0.00	0.16	0.00	0.14	0.00
GP21	0.94	0.00	0.93	0.00	0.76	0.00	0.80	0.00	0.78	0.00	0.77	0.00
GP22	0.93	0.00	0.98	0.00	1.00	0.01	1.00	0.00	1.00	0.00	1.00	0.00
GP23	1.00	0.00	1.00	0.00	0.99	0.00	1.00	0.00	0.89	0.00	0.84	0.00
GP24	0.93	0.00	0.92	0.00	0.82	0.00	0.88	0.00	0.31	0.00	0.21	0.00
GP25	1.00	0.00	1.00	0.00	0.99	0.00	1.00	0.00	0.69	0.00	0.80	0.00
GP28	1.00	0.00	1.00	0.00	0.98	0.00	0.98	0.00	0.95	0.00	0.92	0.00
GP29	1.00	0.00	1.00	0.00	0.99	0.00	0.98	0.00	0.96	0.00	0.90	0.00
GP30	0.99	0.00	0.98	0.00	0.97	0.00	0.97	0.00	0.61	0.00	0.46	0.00
GP31	1.00	0.00	1.00	0.00	0.99	0.00	0.99	0.00	0.73	0.00	0.72	0.00
IP	0.90	0.00	0.94	0.00	0.49	0.00	0.59	0.00	0.08	0.00	0.21	0.00
COSTS	0.96	0.00	0.98	0.00	0.97	0.00	0.97	0.00	0.45	0.00	0.75	0.00
NEER	0.37	0.00	0.56	0.00	0.31	0.00	0.49	0.00	0.52	0.00	0.74	0.00

Note: The table shows the p-values of the null hypothesis of a unit root in the corresponding time series, which are all in logs. Tests are carried out for the full sample. The lag length is automatically chosen. The ADF test is the augmented Dickey-Fuller test and PP test refers to the Phillips-Perron test.

3.2 Empirical framework and results

3.2.1 Long-run vs. short-run exchange pass-through

Our empirical approach starts with a general consideration of the relationship under investigation. Equation (1) provides the basic equation for an ERPT equation (Campa and Goldberg, 2005).

$$p_t = c + \beta_1 s_t + \beta_2 w_t + \beta_3 d_t + \varepsilon_t \quad (1)$$

where p_t corresponds to the index of import prices, c denominates a constant, s_t is the nominal effective exchange rate of the mark which has been converted according to the euro exchange rate after 1999 and d_t denominates German industrial production as a proxy of domestic demand. Finally, w_t denotes a proxy for producers' costs. Owing to a lack of data, we adopt the methodology of Campa and Goldberg (2005), also applied by Mumtaz et al. (2006) and Ben Cheikh (2012a), and use price indices to account for movement in production costs. If a price-setting strategy based on mark-ups over marginal costs is applied, this procedure should capture changes in producers' costs. In addition, price indices are usually highly correlated with wages, which are an important cost factor. Hence, we employ $w_t = \frac{s_t \cdot pp}{r_t}$ as a measure of costs, where r_t is the real effective exchange rate and pp denotes producer prices. All variables enter Equation (1) in logarithms.

In the following, we will refer to the coefficient β_1 as the long-run pass-through coefficient which should, theoretically, take values between 0 and -1. The reason for this is that an increase in the nominal effective exchange rate corresponds to an appreciation in the German currency, which should result in a decrease in import prices. To analyze the short-run, we rely on the following representation in first differences:

$$\Delta p_t = \mathrm{c} + \sum_{i=1}^{n} a_{1,\mathrm{i}} \Delta p_{t-i} + \sum_{i=1}^{m} a_{2,\mathrm{i}} \Delta s_{t-i} + \sum_{i=1}^{o} a_{3,\mathrm{i}} \Delta w_{t-i} + \sum_{i=1}^{p} a_{4,\mathrm{i}} \Delta d_{t-i} + \eta_t \qquad (2)$$

The coefficient $a_{2,1}$ corresponds to the short-run pass-through. The remaining terms on the right-hand side refer to other short-run dynamics and are included to avoid any misspecification due to autocorrelation, while η_t is an i.i.d. variable which follows a normal distribution with zero mean (Enders, 2009). To achieve a distinction between short-term and long-term pass-through coefficients based on Equations (1) and (2), we now proceed by describing the empirical framework in greater detail.

3.2.2 Analysis of long-run pass-through coefficients

Several approaches exist which allow testing for a long-run relationship as given in Equation (1). In this paper, we start with the multivariate approach of Johansen (1988, 1991), which has the main advantage of being completely unrestricted in terms of endogeneity and the number of long-run relationships. Hence, we consider this framework as a starting point, since it enables us to identify the number of long-run relationships and the causality structure of the variables. As will be illustrated below, this is extremely useful when it comes to analyzing the time-varying coefficient models. As a starting point, consider the following vector autoregression representation (VAR) introduced by Johansen (1988):

$$\Delta Y_t = \Pi Y_{t-1} + \Gamma(L)\Delta Y_{t-l} + \Phi D_t + \varepsilon_t \tag{3}$$

The vector Y = (p_t, s_t, w_t, d_t) contains all variables displayed in Equation (1). The long-run level matrices Π can be fragmented into two $r \times p$ matrices α and β′ (Π = αβ′). β′ gives the coefficients of the variables for the r long-run relations, while α contains the adjustment coefficients describing the reaction of each variable to disequilibria from the r long-run relations given by the $r \times 1$ vector $\beta' Y_{t-1}$. The deterministic components are given by the $(p \times 1)$ vector ΦD_t, while ε_t describes an independent and identically distributed error term (Juselius, 2006). The term $\Gamma(L)\Delta Y_{t-l}$ describes the short-run dynamics of the model, which will not be explicitly considered in the following since we consider a time-varying coefficient framework for the short-run at a later stage. To identify the number of cointegrating relations r we rely on the trace test developed by Johansen (1988).

Before we turn to a separate long-run analysis of each price index, it is advisable to consider the subsystem Y\` = $(s_t, w_t, d_t)$. The reason for this is obvious, since the dynamics of Y\` by definition enter any analysis based on Y. From a theoretical point of view, restrictions applied to a subsystem should continue to hold in larger systems (Juselius, 2006). Against this background, an interesting question is whether the number of long-run relationships varies when prices are included in the system. If r is the same for an analysis based on Y\` and Y, no common stochastic trend results when p_t is included and the identification of a long-run relationship in the system under investigation does not depend on p_t. After tackling this issue, we proceed by analyzing the full system Y for each price index.

The findings displayed in Table 3 summarize the results of the rank tests and, where cointegration is detected, provide the estimates of the cointegrating vector and the adjustment coefficient.

Table 3: Results of the multivariate cointegration analysis

	a) p-value	**b) Long-run relationship**	**c) Adjustment coefficient**
Basic system	r=0: 0.110 r=1: 0.469	No cointegration	
GP02	r=0: 0.054 r=1: 0.129	$p_{02} = -1.125^a * s - 2.427^a * w + 0.664^a * d + 0.004^a * t$	-0.151^a
GP03	r=0: 0.040 r=1: 0.209	$p_{03} = 1.170 * s + 10.365^a * w + 5.628^a * d - 0.015^a * t$	0.027^a
GP05	r=0: 0.231 r=1: 0.321	No cointegration	
GP06	r=0: 0.354	No cointegration	

	r=1: 0.528		
GP07	r=0: 0.021 r=1: 0.448	$p_{07} = 1.278 * s + 11.521^a * w + 4.384^a * d - 0.007^a * t$	0.010
GP08	r=0:0.051 r=1:0.694	$p_8 = -0.686 * s + 6.439^a * w + 1.964^a * d + 0.007^a * t$	-0.006[a]
GP11	r=0: 0.274 r=1: 0.971	No cointegration	
GP12	r=0: 0.308 r=1: 0.953	No cointegration	
GP15	r=0: 0.000 r=1: 0.136	$p_{15} = -0.759^a * s + 1.076^a * w + 0.166^a * d + 2.393^a * t$	-0.068[a]
GP16	r=0: 0.001 r=1: 0.053	$p_{16} = 0.148 * s + 1.046^a * w + 0.781^a * d - 0.001^a * t$	-0.021[c]
GP17	r=0: 0.036 r=1: 0.180	$p_{17} = -0.824^b * s - 1.116 * w - 0.849^a * d + 0.003^a * t$	-0.021[a]
GP19	r=0: 0.348 r=1: 0.619	No cointegration	
GP21	r=0: 0.272 r=1: 0.840	No cointegration	
GP22	r=0: 0.002 r=1: 0.150	$p_{22} = -0.020 * s + 2.437^a * w + 0.231^b * d - 0.002^a * t$	-0.018[a]
GP23	r=0: 0.003 r=1: 0.162	$p_{23} = -0.131^b * s + 0.828^a * w + 0.024 * d + 1.295^a * t$	-0.045[a]
GP24	r=0: 0.051 r=1: 0.436	$p_{24} = 1.194^a * s + 3.734^a * w + 2.995^a * d - 0.004^a * t$	0.018
GP25	r=0: 0.000 r=1: 0.079	$p_{25} = -0.105^c * s + 0.884^a * w + 0.265^a * d - 0.202 * t$	-0.050[a]
GP28	r=0: 0.000 r=1: 0.124	$p_{28} = -0.314^a * s + 0.801^a * w + 0.128^a * d + 1.778^a * t$	-0.042[a]
GP29	r=0: 0.064 r=1: 0.784	$p_{29} = -0.010 * s + 1.044^a * w + 0.454^a * d - 0.001^a * t$	-0.009[a]
GP30	r=0: 0.036 r=1: 0.384	$p_{30} = 0.075 * s + 2{,}736^a * w + 0{,}032 * d - 0.001^a * t$	0.020[a]
GP31	r=0: 0.098 r=1: 0.329	$p_{31} = -0.041 * s - 0{,}009 * w - 0.379^a * d - 0.001^a * t$	0.026[b]

Note: Panel (a) provides the p-values of Johansen's (1988) rank test for the null hypothesis that the rank is 0 and 1. Panel (b) shows the estimates of the cointegration vector if cointegration is found. Panel (c) gives the adjustment coefficient for the corresponding import price index towards the long-run equilibrium. [a]/[b]/[c] denote significance at the 1%/5%/10% levels.

No cointegration is detected in the basic system of Y`. However, this changes when import prices are included, since one long-run relationship is not rejected at the 10% level for Y in fifteen out of twenty one cases. In those cases, the existence of two long-run relationships is always rejected.

Identification of the cointegrating system is achieved by normalizing on the corresponding price index. This provides an estimate of Equation (1). An inspection of the long-run estimates for the pass-through coefficient β_1 provides significant coefficients in seven out of fifteen cases. Six of the corresponding coefficients are in line with theory, since the magnitude of the coefficients is between 0 and -1, although forestry and logging (GP02) yield a coefficient slightly smaller. An interesting finding is that the coefficient for metal products (GP24) takes a value of around 1. This finding may be explained by the fact that the corresponding prices do not adjust to long-run deviations, suggesting that dynamics other than pass-through drive the identified long-run relationship. In all other cases where the pass-through coefficient is significant, the adjustment coefficient of prices turns out to be significant and correctly signed, which suggests that the long-run relationship identified mirrors a pass-through mechanism. This conclusion is also supported by the fact that the adjustment coefficient is usually insignificant or wrongly signed in the remaining cases.

Altogether, a theoretically conform pass-through coefficient is only observed in roughly 30% of cases. Thus, an analysis on an aggregated level as, for example, in Kiliç (2010) or Brun-Aguerre et al. (2012) may be biased.

Given that our sample includes several extraordinary events, including the introduction of the euro and the financial crisis, which may alter our results in terms of a long-run relationship, instabilities and structural breaks are important issues. An inspection of recursive maximum likelihood estimates of the cointegrating vectors suggests that the long-run estimates are subject to significant instabilities and the frequently observed insignificance may be traced back to high standard errors in the estimates.[8] De Bandt et al. (2007) and Kiliç (2010) find considerable variation in the long-run pass-through coefficients for Germany, while the disaggregated analysis by Campa and Goldberg (2005) also reveals significant differences across five product categories. For this reason, and owing to the fact that we are also interested in the evolution of the coefficients, we re-estimate the cointegrating relationship for the fifteen categories where cointegration was detected in a time-varying single equation framework with import prices as the dependent variables.

[8] The graphs of the recursive maximum likelihood estimations have been omitted to save space. However, they are available upon request.

Adopting a single-equation approach instead of the multivariate estimator also provides a robustness check and is justified by the fact that a long-run relationship was not detected without import prices. Hargreaves (1994) has shown that single-equation estimators provide efficient estimates.

The framework we use is a state space model which combines the Kalman filter with a regression based on dynamic OLS. More precisely, the basic equations have the following form:

$$p_t = Y'_{t-1}\theta_{k,t} + \varepsilon_t \qquad \varepsilon_t \quad \sim N\left(0, H_{k,t}\right) \tag{4}$$

$$\theta_t = \theta_{t-1} + \eta_t \qquad \eta_t \quad \sim N\left(0, Q_{k,t}\right) \tag{5}$$

Equation (4) is the observation equation and Equation (5) the state equation. θ_t is an $m \times 1$ vector of states corresponding to the coefficients which evolve as random walks. The matrix Q_t corresponds to the variances and covariances of the states and determines changes in the coefficients. The errors in the observation and the state equation covariances are assumed to be mutually independent at all leads and lags. At each point in time Kalman-filtering begins with a prediction of both equations based on an optimization of the projected error covariances. After making a new observation, estimates are then corrected according to the Kalman gain or the blending factor, which minimizes the posterior error covariances. Since we deal with a model based on I(1) variables, the error term may be serially and contemporaneously correlated with the regressors. These assumptions govern the Kalman filter for this model. The DOLS estimator introduced by Stock and Watson (1993) corrects traditional OLS with regard to endogeneity and serial correlation by including leads and lags for the first differences on the right hand side of the equation.[9] In the following, we include two leads and lags for both short-run and long-run pass-through estimations. The results only marginally change in the case of other choices.[10] As a first step, we compare the final estimates of the long-run pass-through coefficients with our previous findings. In order to keep our results clearly laid out, the values of the remaining coefficients, including the leads and lags, are not reported. They are, however, available upon request. The estimates are given in Table 4.

[9] From a theoretical point of view, the OLS estimator is superconsistent in the case of cointegration.
[10] The corresponding results are available upon request.

Table 4: Comparison of long-run pass-through coefficients

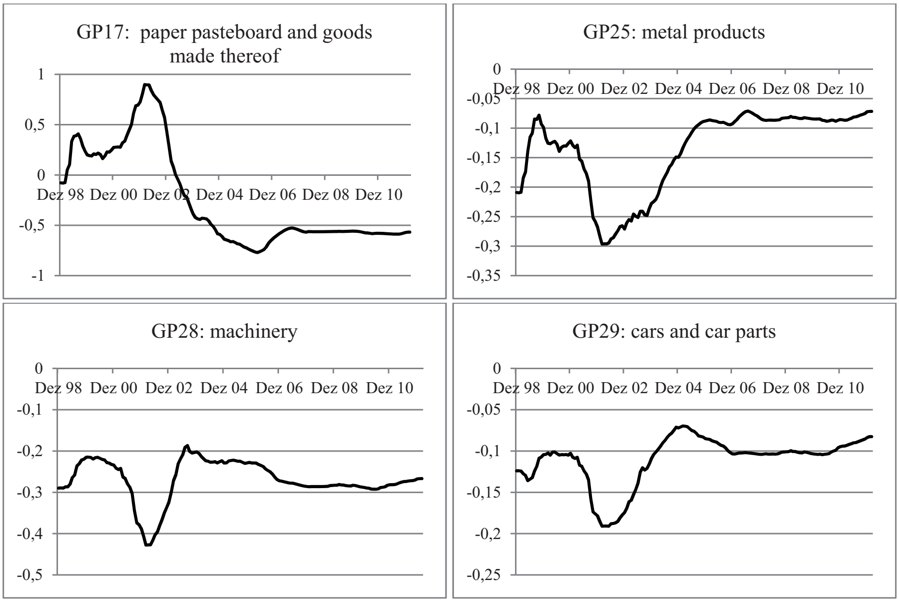
Product category	Johansen	DOLS
GP02	-1.125[a]	-1.000[a]
GP03	1.170	0.837[a]
GP07	1.278	1.265[a]
GP08	-0.686	-0.372[a]
GP15	-0.759[a]	-0.616[a]
GP16	0.148	0.134[a]
GP17	-0.824[b]	-0.568[a]
GP22	-0.020	0.042
GP23	-0.131[b]	-0.079[a]
GP24	1.194[a]	1.090[a]
GP25	-0.105[c]	-0.072[a]
GP28	-0.314[a]	-0.267[a]
GP29	-0.010	-0.083[a]
GP30	0.075	0.033[a]
GP31	-0.041	-0.057[c]

Note: The table shows the long-run pass-through coefficients according to the long-run relationship as shown in Equation (1). [a]/[b]/[c] denote significance at the 1%/5%/10% levels.

We first observe that the results of the full-sample DOLS estimates are very close to the estimates obtained from the multivariate cointegration approach. In all cases, except the category “rubber and plastic goods” (GP22), where both estimates are insignificant, the coefficients have the same sign and roughly the same magnitude. For “stone and other mining products” (GP08), “cars and car parts” (GP29) and “furniture” (GP31), the DOLS estimation provides significant estimates, while the Johansen approach does not. Altogether, the DOLS approach gives a significant pass-through coefficient in fourteen out of fifteen cases. However, five estimates give a reversed and significant sign. The group “metal” (GP 24) had already been observed using the multivariate approach. The other cases include relatively small groups: “mineral ores” (GP07), “fishing and aquaculture” (GP03), “wood, basketry, wattle and cork works (without furniture)” (GP16) and “other vehicles” (GP30).

Based on this finding, Figure 1 provides the evolution of the pass-through coefficients of the four biggest price groups where theory-conform pass-through is detected after the initial estimation period up to 1998:12. The estimates show great variability starting in 2001 but are in line with theory throughout the sample except for “paper, pasteboard and goods made thereof” (GP17), where the coefficient turns out to be positive at the beginning.

Figure 1: Recursive long-run pass-through coefficients



To draw a preliminary conclusion, we have obtained estimates which are remarkably robust in terms of magnitude according to both estimation techniques. The time-varying model provides more frequent results in terms of significance. At the same time, the pass-through differs across product groups. We observe a theoretical pass-through according to both procedures for six out of fifteen groups and nine groups if we only consider time varying estimates. Prices of “metal” seem to be a special case, since the pass-through coefficients show a reversed sign and no price adjustment is observed. In addition, the magnitude of the significant coefficients also differs. Unsurprisingly, pass-through is incomplete in most cases. Bearing these results in mind, we now turn to the time-varying estimates for the short-run pass-through.

3.2.3 Estimations of time-varying short-run pass-through

As our evidence and the literature mentioned in the previous section show considerable instabilities regarding the long-run coefficients, we now neglect cointegration and focus on short-run dynamics. The choice of this procedure is backed up by the fact that one aim of this study is to provide as complete a picture of German ERPT as possible and that the product groups under consideration represent about 60% of German imports. As cointegration is not evident in each category, dropping these groups would narrow the scope of the study and the relevance of the results would be limited.

For the first differences that correspond to short-run pass-through according to Equation (2), we consider two different configurations to achieve time-varying coefficients: a time-varying coefficient model based on the Kalman filter state space approach introduced in the previous section and a recursive estimation via OLS. Both techniques rely on the same idea: after each estimation, a new observation is included, resulting in different estimates at each point in time. However, although they are closely related there are some differences, since the state space model includes stochastic innovations to the coefficients. Altogether, a consideration of both estimators includes an implicit robustness check with regard to our finding. For both approaches, our initial sample period ends in December 1998. This is extended until we estimate our models for the whole sample. This gives us a total of 161 recursive short-run coefficients. The results for both approaches are summarized in Table 5. Furthermore, for the four most important product categories with respect to their share in the total import price index we additionally show the complete evolution of the short-run pass-through coefficients in Figure 2. The detailed results for the other product groups are of course available upon request.

Table 5: Short-run pass-through coefficients

	Initial estimate		**Final estimate**	
	Kalman filter	**OLS**	**Kalman filter**	**OLS**
GP02	-0.16	-0.71[b]	-0.21	-0.40[c]
GP03	-0.76[b]	-0.79[b]	-1.13[c]	-0.99[a]
GP05	-0.69[b]	-0.55[b]	-1.75[a]	-1.35[a]
GP06	-0.82	-0.91	-1.10[b]	-1.18[a]
GP07	-1.78[a]	-1.75[a]	-1.18[a]	-1.33[a]
GP08	-0.34[a]	-0.25[a]	-0.27[a]	-0.19[a]
GP11	-0.16[b]	0.00	-0.01	0.04
GP12	-0.06	-0.12[c]	-0.04	-0.02
GP15	-0.38[a]	-0.23[a]	-0.50[a]	-0.50[a]
GP16	-0.25[a]	-0.33[a]	-0.20[a]	-0.19[a]
GP17	-0.13[c]	-0.17	-0.29[a]	-0.25[a]
GP19	-2.47[a]	-2.55[a]	-2.25[a]	-1.82[a]
GP21	-0.51[a]	-0.42[a]	-0.47[a]	-0.40[a]
GP22	-0.06[b]	-0.05	-0.10[a]	-0.08[a]
GP23	-0.05[c]	-0.05	-0.06[b]	-0.03[c]
GP24	-1.03[a]	-0.96[a]	-0.74[a]	-0.43[a]

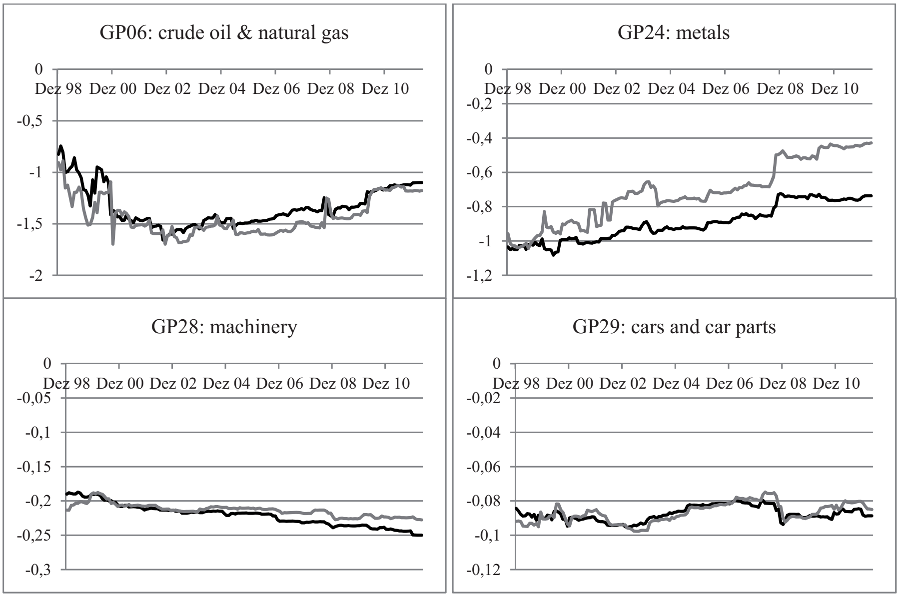
GP25	-0.20[a]	-0.19[a]	-0.21[a]	-0.19[a]
GP28	-0.19[a]	-0.21[a]	-0.25[a]	-0.23[a]
GP29	-0.08[a]	-0.09[a]	-0.09[a]	-0.09[a]
GP30	-0.07	-0.08	-0.23[a]	-0.20[a]
GP31	-0.08[a]	-0.07[b]	-0.15[a]	-0.14[a]

Note: The table shows initial and final estimates of the short-run pass-through coefficients $a_{2,1}$ according to Equation (2). The initial estimate refers to the sample up to December 1998, while the final estimate considers the whole sample up to April 2012. [a]/[b]/[c] denote significance at the 1%/5%/10% levels.

For our time-varying coefficient model, we generally observe negative short-run pass-through coefficients which are significantly different from zero in nearly all cases. Additionally, the results obtained via both OLS and the Kalman filter yield very similar results. The evolution as well as the magnitude of the coefficients is quite similar in both approaches. Differences are probably not statistically significant. Therefore, the results seem to be remarkably robust. Hence, short-run pass-through shows a more uniform character compared to the long-run pass-through estimates. An inspection of Figure 2 shows that the coefficients are less volatile compared to the time varying long-run coefficients displayed in Table 1.

Nevertheless, the evolution of the coefficients differs across product groups. There are cases where the short-run pass-through is stable over time (mineral ores (GP07), stone and other mining products (GP08), pharmaceutical goods (GP21) and metal products (GP25)), where it increases i.e. the coefficient becomes more negative (coal and peat (GP05), leather and leatherware (GP15), paper, pasteboard and goods made thereof (GP17), other vehicles (GP30), furniture (GP31)) and where it decreases (forestry and logging (GP02), beverages (GP11), tobacco products (GP12), metals (GP24)).

Figure 2: Recursive short-run pass-through coefficients



Note: The estimates based on the Kalman filter are represented by the black line. The grey line stands for the OLS estimates. The date corresponds to the end of the estimation sample.

For "mineral ores (GP07)", where the long-run coefficient showed a reversed sign, we now observe a coefficient which is even smaller than -1. Moreover, this over-proportional short-run pass-through seems to be stable over time. The short-run pass-through coefficient for "metals", where we also observed an incorrectly signed long-run coefficient, provides an estimate of about -1 at the beginning of the sample. Over time, its magnitude decreases. Regarding the magnitude of the coefficient estimates of product category 19 (cookery and mineral oil products), the short-run pass-through is around 2.5, which implies a very strong reaction of import prices to exchange rate movements.

Regarding imported inflationary effects, it seems that short-run pressures decreased (or at least did not increase) over time. For products with large shares in the overall import price index (see Figure (1), i.e. crude oil and natural gas, pharmaceutical goods, metals, machinery, cars and car parts) , the pass-through either decreases or remains fairly stable. Such a pattern is also observed for Germany by Brun-Aguerre et al. (2012) on an aggregated level. Mumtaz et al. (2006) find decreasing pass-through rates for disaggregated UK import prices. A considerable increase in the short-run pass-through coefficient is observed only for product categories of minor importance (coal and peat, other vehicles, furniture). A comparison of our results to other studies is somewhat difficult. Only a few investigations use disaggregated data and the disaggregation is generally not as broad as ours. For example, Campa and Goldberg (2005) estimate pass-through elasticities for five different product categories. According to their results, products that are less processed exhibit larger pass-through. Such a picture, which is a general finding in the literature,[11] is evident in our sample too. For example, highly processed goods such as "machinery (GP28)", "cars and car parts (GP29)" and "other vehicles (GP30)" show small short-run pass-through coefficients. In contrast, the coefficients for "crude oil and natural gas (GP06)", "mineral ores (GP07)" and "metals (GP24)" are close to one or even above full pass-through.

3.2.4 What drives exchange rate pass-through?

Having analyzed the degree and character of pass-through for these different product categories, we now turn to our second research question and analyze the evolution of pass-through against the background of different macroeconomic factors. Taking into account the previous findings, we now concentrate on the short-run pass-through. The fact that long-run pass-through has only been detected in at most nine cases and rejected for some large groups such as metals makes the analysis of common driving factors more sensible for the short-run. Since we are interested in the question of whether a similar pattern across different groups can be observed, we tackle this question in a panel framework. In order to do this, we estimate our pass-through model (Equation (2)) for each of our twenty one product categories for rolling windows.[12] For convenience, we perform this analysis only via OLS, as the previous section has shown that the differences in OLS and Kalman filter models are negligible. In doing this, we obtain twenty one series of short-run pass-through coefficients. These coefficients are then used as endogenous variables in a second-step regression on various macro-determinants. To be more precise, when pooling the data we consider the following specification:

$$pt_{it} = c + macro_{it} \cdot \beta + \varepsilon_t. \tag{6}$$

pt_{it} represents the short-run pass-through coefficients obtained from the rolling regressions (Equation (2)), c is a constant term and $macro_{it}$ is a matrix of the factors considered as determinants of the pass-through. These factors are explained in more detail in the following paragraph.

In addition to the pooled estimation we estimate a fixed-effects model, which is represented by the following equation:

[11] Al-Abri and Goodwin (2009) state that the fact that these goods are homogenous and that their prices are quoted in US dollars allows a fast adjustment of local prices.
[12] We use rolling estimates because these will probably exhibit changes in coefficients more quickly than recursive estimates.

$$pt_{it} = \alpha_i + macro_{it} \cdot \beta + \varepsilon_{it} \quad (7)$$

The difference from the pooled OLS specification is that we now assume specific intercept terms for each product category.

Determinants of the short-run exchange rate pass-through

We consider the following variables as determinants of the exchange rate pass-through that are typically used in the literature (Mumtaz et al., 2006, Brun-Aguerre et al., 2012). All determinants enter equations as averages over the estimation period.

Inflation: For inflation, we use both import (INFL_AVG) and consumer price inflation (CPI_AVG), which is measured as the year-on-year change in the corresponding price index. Accordingly, we have a product-specific inflation rate (i.e. the import price inflation) and the broader CPI inflation rate. We incorporate both inflation rates because these can deviate considerably from each other. As Taylor (2000) has outlined, higher inflation rates are associated with a stronger pass-through, which corresponds in our case to a negative relationship. Hence, higher inflation rates should result in a more negative pass-through coefficient.

Inflation volatility: As for inflation, we consider both the volatility of import (INFL_VAR) and consumer price inflation (CPI_VAR). This volatility is computed as the standard deviation in the inflation rate over the estimation period.

Exchange rate: As for inflation rates, we consider both the average (NEER_AVG) and the volatility of the exchange rate (NEER_VAR) as explanatory variables. We use the same exchange rate series as in the foregoing analysis, namely the nominal effective exchange rate. From a theoretical point of view, it is not clear whether there is a positive or a negative link. On the one hand, when firms try to maintain market shares, they will adapt their mark-ups and high volatility is associated with a lower pass-through (Froot and Klemperer, 1989). On the other hand, if exporters are particularly interested in their margins, exchange rate changes should be transmitted to import prices rather quickly and completely, indicating a positive link.

Exchange rate surprise: Similar to Mumtaz et al. (2006), we consider forecast errors in the nominal and real effective exchange rate (FE_NEER, FE_REER). Unexpected changes in the exchange rate might either increase or decrease the pass-through according to whether enterprises try to maintain market shares or keep margins constant. These unexpected exchange rate changes are calculated as forecast errors in a simple three-month autoregressive exchange rate forecast.[13]

Output gap: A positive output gap might be associated with a lower exchange rate pass-through if firms squeeze margins to satisfy the rising domestic demand for foreign products. We calculate the output gap as a deviation in German industrial production (in logs) from its Hodrick Prescott filtered trend. As with the other variables, we consider the average (GAP_AVG) as well as the volatility (GAP_VAR) of this variable.

Table 6 shows the results of the regression of the rolling short-run exchange rate pass-through coefficients on its supposed determinants for, respectively, the fixed-effects and pooled-sample specifications. To ascertain the determinants we followed a general-to-specific approach. As a first step we included all the aforementioned determinants in the regression and gradually dropped the least significant regressor until we were left with coefficients which were significant at least at the 10% level.

Table 6: Determinants of short-run ERPT

	10y window		8y window		5y window	
	FE	**Pooled**	**FE**	**pooled**	**FE**	**pooled**
Constant		4.58[b]		0.42[a]		-0.05[a]
NEER_AVG	-0.03[b]	-0.05[b]			12.50[a]	7.61[a]
NEER_VAR	-0.03[b]	-0.04[c]	-0.15[a]	-0.14[a]	0.03[b]	
INFL_AVG	-0.02[a]	-0.01[a]	-0.02[a]	-0.02[a]		-0.01[a]
INFL_VAR	-0.06[a]	-0.05[a]	-0.04[a]	-0.04[a]	-0.04[a]	-0.04[a]
GAP_AVG	3.03[b]		4.98[a]	4.68[b]		
GAP_VAR			-3.59[a]	-3.64[b]		
CPI_AVG					-0.21[a]	
CPI_VAR	0.32[c]	0.52[c]				
FE_NEER			-0.70[a]	-0.65[a]		0.34[a]
FE_REER			0.21[a]	0.19[b]	-0.21[a]	-0.10[a]
Observations	1617	1617	2121	2121	2877	2877

Note: [a]/[b]/[c] denote significance at the 1%/5%/10% levels. The columns labeled "FE" gives the coefficients of the fixed-effects specification according to Equation (7). The columns entitled "pooled" show the coefficients of the pooled sample estimation (Equation (6)).

As can be seen from Table 6, there is no factor which shows up in no specification. Accordingly, all determinants seem to influence the pass-through at least in some cases.

[13] Forecasting exchange rate is a notoriously difficult task. We do not consider more sophisticated techniques, partly because no approach is able to systematically beat a simple random walk forecast, as first outlined by Meese and Rogoff (1983).

However, the coefficients sometimes differ considerably across the models. But, nevertheless, some results do seem to be robust.

For example, consider the coefficient of the volatility of the nominal effective exchange rate (NEER_VAR). This coefficient can be found in five out of six models and is negative in four cases, indicating that higher exchange rate uncertainty is associated with more negative pass-through coefficients. This finding suggests that firms do not hesitate to pass through exchange rate movements but rather feel free to adjust prices quickly. This might indicate market power. The next variable that seems to be a driver of exchange rate pass-through is the average product-specific inflation rate (INFL_AVG). It turns out to be significant in five models. Although its magnitude is always close to zero, it is negative and highly significant in each case. Higher inflation rates correspond to higher (more negative) pass-through coefficients. Such a finding is supported by Kiliç (2010). The sector-specific inflation volatility (INFL_VAR) enters as significantly negative in each of the six specifications. As for the exchange rate, uncertainty increases the amount of short-run pass-through. The average output gap is positively linked to the degree of short-run ERPT. If the German economy is running above potential this implies a muted pass-through. For all other determinants we have considered, the evidence is either weak or inconclusive. Estimates for consumer price inflation variability or average only turn out to be significant in three cases which suggest that import prices are more important for the evolution of the pass-through mechanism. Effective exchange rate forecast errors turn out to be significant in three (nominal effective exchange rates) and four (real effective exchange rates). However, sign and magnitude of the coefficients differs which is not surprising considering that theoretical suggestions are ambiguous. From a general perspective, variability provides more theory conform estimates compared to averages, suggesting that changes in degree of pass-through may be traced back to uncertainty over macroeconomic fundamentals in terms of high volatility.

4. Conclusion

Putting the exchange rate pass-through mechanism into German import prices for disaggregated data under closer scrutiny, we have illustrated the complexity of the underlying dynamics. Firstly, a distinction between short- and long-run pass-through is necessary. While pass-through is only partly observable and incomplete for the former, short-run pass-through shows a more unique character, although heterogeneity across product groups still exists. Another finding is that the degree of pass-through is subject to several changes. Applying a variety of recursive estimation techniques, we have illustrated that the short-run pass-though coefficients do change over time but mostly move within a certain range. For the main categories, the short-run pass-through remains fairly stable or decreases slightly. By contrast, the observed long-run pass-through coefficients show much more variability, which is in line with previous research. Hence, empirical results on long-run pass-through crucially depend on the sample under investigation, with data mining being a serious issue.

Finally, several macroeconomic factors are responsible for the observed changes in the degree of short-run pass-through. The results are partly ambiguous in terms of different time windows but price and exchange rate volatility turn out to be frequently important factors. Hence, the degree of pass-through is higher during periods of high volatility, possibly due to a premium as a compensation for the resulting uncertainty.

Altogether, several issues for further research remain. Besides analyzing a longer span of data, a comparison between Germany and other economies would be an interesting task. Furthermore, allowing for nonlinearities might be another extension. From a policymaker's perspective, our results provide important insights. In general, there does not seem to be increasing pressure on inflation rates via the import price channel, which is relevant for open economies like Germany. However, a disaggregated view of pass-through is reasonable because considerable heterogeneity, especially with regard to the magnitude of ERPT across product categories, does exist. Moreover, volatile exchange rates, as in the years of the financial crisis, or increasing inflation rates and inflation volatility due to the huge liquidity provided by central banks might reinforce pass-through. Analyzing the implications of our findings for exchange rate adjustment in the context of global current account imbalances as outlined by Obstfeld and Rogoff (2005) is another interesting topic for further research.

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