The Empirical Range of Pass-Through in US, German and Japanese Macrodata

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Abstract

This paper compares the empirical range of aggregate exchange rate pass-through in the US, Germany and Japan during the 1980s and the1990s. Our main contribution is to focus on monthly data, to better account for real-world price level stickiness and exchange rate volatility. Another import is that we take robustness seriously and obtain our results employing a battery of alternative specifications, including notably generalized VARs. We find that, first, pass-through on import prices has considerably declined in the 1990s relative to the 1980s, pass-through on export prices has not changed much and pass-through on consumer prices seems to be nowadays practically negligible in all three countries we considered. Second, the econometric method, variable proxy and data frequency used matter for the precise magnitudes and time patterns, yet they often accord on the general trends; our emphasis on monthly series has, however, uncovered a common profile of short-run pass-through dynamics which remains hidden in quarterly observations. Third, the US is quite a particular economy, with import and, hence, consumer price levels that are amazingly insensitive to US dollar fluctuations.

JEL Classification: F10, F33, F41.

Keywords: exchange rate pass-through, international comparisons, sensitivity analysis, generalized VAR estimation.

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1 Motivation, Objective and Approach

An important issue in international economics with consequences for the transmission of shocks across countries that is currently being reconsidered is in what currencies monopolistically competitive firms preset the prices for the products they sell in global markets. In particular, various papers of the new open-economy macroeconomics (NOEM) literature[1] have pointed out in microfounded general equilibrium set-ups that the assumption of consumer's currency pricing (CCP)[2] versus producer's currency pricing (PCP)[3] is of an essential nature under price rigidity. Extending this research to a purposeful parallel of a CCP to a PCP model version, the stochastic NOEM framework in Mihailov (2003 a, b) has analyzed theoretically this alternative invoicing possible in open economies[4] under the extreme assumptions that either CCP is full for both interacting economies or PCP is, in turn, complete. Although rather simplistic, such a modelling strategy has helped clarify why from an economy-wide viewpoint the currency in which prices are preannounced is crucial in equilibrium trade and exchange rate determination. The reason is that *full* CCP – by preventing any pass-through from nominal exchange rate changes to import and, ultimately, consumer prices – completely reverses a central result in the Keynesian international macroeconomics tradition known as the *expenditure-switching effect*: a monetary expansion that depreciates the national currency – and hence, within the short run of price stickiness, the real exchange rate – leads under full CCP to an improvement (not deterioration, as under full PCP) in the inflating country's terms of trade (ToT) and ultimately depresses (and does not stimulate) real economic activity.

In reality, however, CCP and PCP will coexist in the prices of exported as well as imported products, and the *extent* of CCP (or, inversely, PCP) would thus largely determine the empirical range of pass-through from nominal exchange rate changes to import, producer, consumer and export prices of a given country. In Mihailov (2003 a, b) we have also shown that pass-through will be a key factor – together with the additional influence of transport, tariff and related frictions and of the elasticity of substitution between the good-types produced by the two economies – in accounting for any trade growth and stabilization role of the exchange-rate regime. We have concluded from our theoretical analysis that, with monetary uncertainty and nominal rigidity, more CCP in bilateral trade invoicing would mean less pass-through from the exchange rate to

[1] As defined and classified in the recent survey by Lane (2001). A narrower and more technical summary of the basic NOEM methodology is also provided in Sarno (2001).

[2] Also termed pricing-to-market (PTM) or local currency pricing (LCP).

[3] The standard assumption of open-economy models, both theoretical and empirical, in the Mundell-Fleming-Dornbusch tradition.

[4] Friberg (1998) argues that the currency of *price setting*, the currency of *invoicing* and the currency of *payment*, although theoretically corresponding to three distinct stages of a typical international trade transaction and hence potentially different, practically coincide "with few known exceptions". Therefore we use "invoicing" and "price setting" interchangeably in the present work (without talking at all about the "currency of payment").

import and consumer prices, so a peg would not achieve much in stabilizing national trade shares in output, neither – under inelastic demand for cross-country output – in increasing expected trade. On the contrary, if PCP is the dominant trade pricing convention between two (symmetric) economies so that the degree of pass-through is huge and induces considerable expenditure switching, fixing the exchange rate would always lead to trade stabilization, and also to some trade growth under inelastic import demand.

The objective of the present paper is to further examine empirically the unresolved issue of what is the likely *range* of aggregate exchange rate pass-through. One approach to do this would be to rely on survey data and study the *direct* evidence on currency denomination in actual international trade transactions. Many papers did pursue such an approach in the late 1970s and early 1980s, to generally find that trade in manufacturing goods between developed countries was mainly invoiced in accordance with PCP. This regularity has been referred to as "Grassman's law", after the important empirical contributions by Grassman (1973 a, b) based on 1968 Swedish trade data. Similar applied work, but using more recent (post Bretton-Woods) data, such as Friberg and Vredin (1997), for example, has however supported an increasing role of pricing-to-market (that is, CCP) practices.

An alternative strategy to study the range of pass-through is the *indirect* one, which exploits pertinent data and theoretically postulated relationships underlying their structure and/or dynamics to estimate and interpret key correlation and regression coefficients (elasticities).[5] Following this latter approach, we are interested here in extracting from macroeconomic time series robust interval estimates of pass-through in the three countries whose currencies have been the major international medium of exchange and store of value over the last half of a century, namely the United States (US), Germany and Japan.[6] Similarly to some of the previous literature, we measure exchange rate pass-through at three stages, i.e. on import, export and consumer prices. Yet a particular feature of our analysis which distinguishes it from preceding ones is that we purposefully focus on *monthly* data, this frequency being more relevant to price rigidity predominating in the real world. Another novelty in pass-through research we introduce with this paper is that apart from comparing our results (i) across the three largest national economies nowadays and (ii) across stages along the pricing chain, we essentially perform an exhaustive sensitivity analysis across four additional dimensions: (iii) frequency, (iv) time, (v) econometric methods and (vi) aggregate import/export price proxies and business cycle controls.

The frequency dimension of the empirical analysis relates our findings based on monthly data to their analogues obtained from the same estimation but with *quarterly* data. The time dimension – in effect, an indirect test of Grassman's law with recent data – consists in splitting up the whole sample in two symmetric *subperiods*, the 1980s and the 1990s, to look into the dynamic characteristics of the phenomenon. The methodology dimension of our approach progressively interprets (a) ToT-exchange rate *correlations* as in Obstfeld and Rogoff (2000), (b) *ordinary least squares (OLS)* regressions as in Campa and L.Goldberg (2002) and (c) *vector autoregressions (VARs)*, applying *orthogonalized* impulse responses as in McCarthy (2000) and Choudhri, Faruquee and Hakura (2002) as well as – innovatively in the present study – *generalized* impulse responses, first proposed by Pesaran and Shin (1998), where ordering does not matter. Moreover,

[5] An extensive and widely cited (but now somewhat old) survey of the empirical pass-through literature is P.Goldberg and Knetter (1997).

[6] For instance, in April 1998 the average daily foreign exchange market turnover has been estimated by the Bank for International Settlements (1999) – Statistical Annex, Table E-1 – to be 1260 billions of US dollars in the United States, 430 billions of US dollars in Germany and 300 billions of US dollars in Japan. United Kingdom comes next, with 157 billions of US dollars, followed by Switzerland with 101 billions of US dollars and France with 73 billions of US dollars.

we perform a *battery* of seasonality and stationarity tests and report in explicit detail the conclusions from them, something rarely done in the literature unless in a footnote or two. We also carefully test for Granger causality and *cointegrating relations* suggested by theory to check for possible use of cointegrated VAR models, as recently done by Coricelli, Jazbec and Masten (2003). A final comparison is effected along the proxy dimension, with alternative proxies employed for both trade prices and business cycle indicators: we parallel estimates obtained using the more relevant aggregate import and export *price indexes* with corresponding ones based on the more readily available approximations of the mentioned indexes which are the *unit values* of imports and exports;[7] furthermore, we check how *industrial production* and *employment volume* indexes affect the magnitudes of pass-through when replacing *real gross domestic product (GDP)* as a standard business cycle control variable.[8] In fact, our different measurement strategies to appropriately quantify pass-through build upon one another in a complementary way, correcting for weaknesses in each one of them if applied in isolation.

All papers we quoted – except Coricelli, Jazbec and Masten (2003) whose data are anyway limited to 1993-2002 and four European Union (EU) candidate economies – have relied on quarterly time series, mainly due to lack of monthly import and export price indexes on a wide and comparable international basis. Most of these authors, including McCarthy (2000) and Choudhri, Faruquee and Hakura (2002), have however admitted that monthly data would be more desirable in studying this particular issue, as providing a better approximation to documented rigidity characterizing real-world prices. We therefore exploit essentially this line of empirical inquiry, hoping to improve on earlier pass-through estimates due to the use of data at a higher, and more relevant, frequency as well as to complement them by a thorough sensitivity analysis. Our approach of focusing on monthly time series becomes possible, it is true, within a rather narrow cross-section to ensure highest comparability.[9] Yet the initial country sample here can subsequently be extended to other economies, in particular when both price indexes and unit values of imports and exports are available for long (and coinciding) time spans, an extremely rare feature in the currently available national macroeconomic accounts.

Our results have confirmed that the use of monthly data is quite central when it comes to measuring pass-through more precisely. This is not surprising, since pass-through has to do with reactions of monopolistically competitive price-setters to (i) exchange rate movements (ii) under sticky prices. On both counts, quarterly observations would miss much of the "action". The New Keynesian literature has now converged to a broad agreement that the dynamics of real-world price rigidity, itself often narrowly related to exchange rate volatility, and, hence, the resulting pass-through, is usually better observed at a frequency lower than one quarter. Accordingly, we establish that quarterly data tend to *underestimate* pass-through and to somewhat *distort* its time profile when compared to corresponding monthly based ones, due to certain averaging out of shorter-run price adjustments to changes in exchange rates. Moreover, insofar most previous pass-through estimates have depended on quarterly data, we would claim that our present contribution has improved on earlier quantification in terms of both precision and robustness. Precision, because the monthly frequency matters indeed

[7] It should be noted that the importance of such a dimension of the study originates in the difference in the method of calculating these two trade price proxies. Whereas indices are computed via direct (but not systematic) surveys of exporters and importers concerning the actual prices of international trade transactions, unit values are indirectly obtained from customs declarations registering both volumes and values by transaction. Unit values are, therefore, less reliable although more easily available on a broad basis.

[8] It might also be interesting to try some *output gap* measure in addition to the three aggregate demand proxies enumerated. Yet calculation of output gaps on an internationally comparable basis is rather problematic methodologically and may thus introduce more noise into the estimates.

[9] And at the cost of interpolating GDP series and related deflators in those econometric specifications which include real GDP.

when measuring pass-through, as we have just stressed. Robustness, firstly, because of the three times higher number of observations provided by a monthly sampling relative to a quarterly one within the same period; and, secondly, because we have come up with sort of "interval" estimates for the empirical range of pass-through from pooling together magnitudes obtained by a variety of complementary econometric techniques and variable proxies.

To summarize our conclusions in a preview, we find that the empirical range of exchange rate pass-through varies across (i) countries, (ii) data frequencies, (iii) time periods, (iv) econometric methods, (v) aggregate price and volume proxies, (vi) stages along the pricing chain (import, export and consumer prices) and (vii) time horizons (one month, one quarter, one year). Any generalization should, in consequence, be done carefully and to the extent particular cases lend themselves to it. Leaving aside the specificity concerning some aspects of our pass-through quantification, which we shall discuss in detail further down, we could emphasize here at least three important and quite robust results. First, in the economies we focus on, pass-through on import prices has considerably declined in the 1990s relative to the 1980s but pass-through on export prices has not changed much; as to consumer prices, pass-through has always been practically negligible over all horizons of up to one year. Grassman's law seems thus weaker nowadays as compared to the Bretton-Woods era. Second, the econometric methods and the measurement proxies we used do matter (more so for our price proxies, less so for our volume proxies) for the *precise* magnitudes and time patterns, yet they agree on the *general* tendencies. Third, the US is an economy with import price levels that are astonishingly irresponsive to nominal exchange rate changes, as has also been found in other pass-through studies.

The paper is further down organized as follows. Section 2 describes the data, reports the results from our seasonality and stationarity tests and presents correlations between the terms of trade and the nominal exchange rate, the latter being indicative of predominance or not of CCP vs. PCP in foreign trade invoicing. The third section then discusses the most common approaches to estimating pass-through in related research and motivates our own empirical strategy. Section 4 interprets our estimates across the several dimensions of the present analysis at each stage of the underlying pricing chain, and the fifth section concludes. Definitions of the data, graphical illustrations and descriptive statistics are provided in Appendix A, while Appendix B documents in detail the results from our econometric work.

Throughout the paper, we present and comment our pass-through estimates based on *monthly* data. The corresponding quarterly based estimates are thus only mentioned for comparison purposes and to reveal the differences – at times considerable in magnitude but less so in time pattern – we detect from the same underlying series across the frequency dimension.

2 Data and Preliminary Tests

Our sample is largely based on International Monetary Fund (IMF) data downloaded from the online version of *International Financial Statistics (IFS)* accessible via Datastream. As nominal GDP and GDP deflators are released in quarterly frequency, they were first *interpolated* by the spline method and the corresponding real GDP was then included as a control variable in some of our monthly specifications. An additional data source, in particular for the employment volume index, is *Main Economic Indicators (MEI)* published by the Organization for Economic Cooperation and Development (OECD) and downloadable via Datastream again. Since a monthly series was not available for Germany, estimations for this particular country based on the employment volume index as an alternative business cycle indicator were effected only at the quarterly frequency. The definitions of all data we use here and their respective unique IFS or MEI codes are provided in Appendix A.

2.1 Descriptive Statistics

To obtain higher comparability, we worked on purpose with a sample period divided in two *equal* halves that is completely identical for all our three economies. To circumvent a discontinuity in the IFS money supply series for Germany, which changed in January 1999 the unit of measure from deutsche marks (DEM) to euros (EUR), the German M1 aggregate was expressed in marks for the *entire* sample period.[10] Thus, our whole sample contains 276 monthly observations (1979:07 – 2002:06), with each of the two subsamples, "the 1980s" (1979:07 – 1990:12) and "the 1990s" (1991:01 – 2002:06) covering 138 observations. Graphs (in natural-logarithm levels) and descriptive statistics (in percentage changes) of the monthly series entering our principal specifications for estimating pass-through are provided in Appendix A.[11]

2.2 Testing for Seasonality

The national sources of the data reflected in the original Datastream series are quite heterogeneous, and not all of these variables had been systematically treated for seasonality. To deal with this problem, we relied on explicit seasonality tests by performing the Census X12 procedure. To conclude whether a series displays a seasonal pattern or not, we looked at four formal tests within Census X12. If at least three of the tests indicated presence of some form of seasonality, we considered the time series in question seasonal and further used in our estimation the corresponding deseasonalized variable (again produced via Census X12). In the rare cases where two of the Census X12 tests have indicated seasonality whereas the other two not, we attributed the decisive weight to the *combined* test for identifiable seasonality. Our seasonality test results are summarized in Table 1 in Appendix B.[12]

2.3 Testing for Stationarity

We applied a similar test-diversified procedure when deciding on stationarity issues related to the time series involved in our pass-through estimation. More precisely, we employed three tests that methodologically complement one another, with each of them having been effected in four alternative specifications. Augmented Dickey-Fuller (ADF) unit root tests based on autoregressive models were thus performed in parallel with kernel-based Phillips-Perron (PP) unit root tests, with the null for both tests being that of a unit root (i.e. nonstationarity) present. These two tests were further supplemented by a test constructed on the opposite null, of stationarity, namely the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test, and both autoregressive and kernel-based specifications of it were used. The results from our four specifications of each of the (non)stationarity tests are summarized in Table 2 in Appendix B.[13] Our conclusion whether a time series is stationary (i.e. integrated of order 0, I(0)) or not and whether, if nonstationary, it is integrated of order 1 or 2 (I(1) or I(2)) was based on this latter table. In many cases our three tests have agreed quite unanimously on the order of

[10] We converted the post-EMU EUR-denominated data into DEM-denominated equivalent applying the exchange rate of 1.95583, which was the same on 31 December 1998 and on 1 January 1999.

[11] In addition to our *evenly*-split sample, we have also performed identical monthly-data estimations with an alternative, *asymmetric* split of our sample. It is based on Chow breakpoint and forecast tests for parameter stability and will be explained in more detail in due course. Such a check for robustness has not essentially changed the main tendencies in pass-through we report further down.

[12] Further details, including the EViews programs, are available upon request.

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integration to be 1. What is worth pointing out here – in particular with respect to Coricelli, Jazbec and Masten (2003) who have argued that most prices in transition economies seem to be integrated of order 2 and have consequently used a cointegrated I(2) VAR in deriving impulse responses to measure pass-through – is that we do not find (overwhelming) evidence of I(2) series in our data. Their claim is thus perhaps either a characteristic of transition economies which cannot be generalized or an artefact of the somewhat short sample they use (monthly data over 1993-2002).

2.4 ToT-NEER Correlation Analysis

In a preliminary look into the determinants of pass-through, we now refer to our theoretical results in Mihailov (2003 a, b) concerning the microfounded definition of the terms of trade (ToT) under CCP vs. PCP. We use this definition in a way suggested by Obstfeld and Rogoff (2000) to motivate and replicate their ToT-NEER quarterly correlation tests for the empirical prevalence of one of these types of price-setting, within our sample and at the more relevant monthly frequency. As explained in the beginning, a (high) positive ToT-NEER correlation evidenced in the data may partly be due to a (strong) prevalence of traditional PCP in the foreign trade of a given country. Conversely, a negative correlation or an approximate absence of correlation would signal, among other things, a much greater importance (if not dominance) of CCP behavior consistent with pricing-to-market arguments. Our correlation findings are presented in Table 3 in Appendix B, by country and by subperiod.

The monthly and quarterly ToT-NEER correlations we have computed[14] are practically the same, and are both very sensitive to the time period over which they are measured. For our whole sample, which is the period that compares most directly (although not exactly) with that of Obstfeld and Rogoff (2000), our correlations are much lower than the respective quarterly ones documented by the latter two authors for the US but much higher for Germany and Japan. Furthermore, these correlations fall in the 1990s relative to the 1980s, weakly for Germany and drastically for Japan, but slightly increase in the US. This would suggest a falling degree of pass-through, partly due to an increasing portion of CCP in trade transactions, in Germany and Japan but a reverse tendency, although weak, in US trade prices. We return to these observations in a more careful pass-through analysis and with some possible explanations in sections 3 and 4.

3 Pass-Through Estimates

We next discuss alternative econometric methods of extracting estimates of pass-through from macroeconomic data that have recently been used in influential, or at least widely referred to, papers. In doing so, we also relate our approach to those in the quoted literature. In our empirical analysis further down, we are particularly interested to compare *single-equation* with *system* estimates of pass-through, i.e. the two model specification strategies usually suggested thus far. That is why we highlight the mentioned strategies in the two respective subsections below and then report our findings when employing each one of them in section 4.

3.1 Single-Equation Pass-Through Estimates

When it comes to single-equation pass-through estimates, the most recent study – also summarizing the preceding literature and trying to improve on it – is Campa and

[14] For log-levels of the respective variables, following Obstfeld and Rogoff (2000).

L.Goldberg (2002). For this reason, we first apply their OLS methodology to our sample and two subsamples by country and compare our *monthly*-based estimates on pass-through on *import* prices with the respective *quarterly* ones we also calculate, as well as with the *quarterly* estimates in the cited paper. Several specifications, starting from the original one in Campa-L.Goldberg (2002), were used to infer our pass-through measures. For comparability purposes, we report results only from the model which corresponds exactly to that in Campa and L.Goldberg (2002), but adjusted to account for the change from quarterly to monthly data in the lag structure and for the autocorrelation found and corrected for in the residuals.[15]

Tables 4 and 5 in Appendix B document our results for the three countries, two subperiods and two aggregate price proxies from our principal OLS regression, which is the following:

$$d\ln\left(PMI_{i,t}\right) = c_0 + \sum_{k=0}^{12} c_{1,k} d\ln\left(NEERInv_{i,t-k}\right) + \sum_{k=0}^{12} c_{2,k} d\ln\left(CGCost_{i,t-k}\right) +$$

$$+ c_3 d\ln\left(GDPR_{i,t}\right) + \sum_{k=1}^{3} c_{4,k}\left(AR_{i,t-k}\right) + u_{i,t}, \quad (1)$$

where $PMI_{i,t}$ is the import price index for country i at time t; $NEERInv_{i,t}$ is the nominal effective exchange rate (NEER) index defined inversely to the IFS-Datastream original series to correspond to the usual interpretation of depreciation being the increase (not decrease) in the exchange rate, with k in (1) indexing the time lag; $CGCost_{i,t} \equiv \frac{NEER_{i,t}CPI_{i,t}}{REER_{i,t}}$ is a measure of overall competitiveness Campa and L.Goldberg (2002) suggest as a key control variable, with $CPI_{i,t}$ being the consumer price index (CPI) and $NEER_{i,t}$ and $REER_{i,t}$ being respectively the nominal (NEER) and real (REER) effective exchange rate indexes as defined in IFS-Datastream;[16] $GDPR_{i,t}$ is real GDP;[17] $AR_{i,t}$ are autoregressive error terms added to correct for identified serial correlation in the disturbance process, most likely of order 1, 2 or 3 (according to Durbin-Watson tests and Breusch-Godfrey Lagrange multiplier tests we performed); $u_{i,t}$ is the error term.

In order to judge about the effect of employing alternative aggregate import price proxies, in the cases of Germany and Japan (but not for the US, due to lack of data) equation (1) was also estimated with $PMU_{i,t}$, the unit value of imports, replacing $PMI_{i,t}$ above. Furthermore, we applied other business cycle proxies as controls reflecting aggregate demand conditions, and available at a monthly frequency: firstly, we replaced $GDPR_{i,t}$ by $IPI_{i,t}$, the industrial production index; secondly, in the cases of the US and Japan (but not for Germany, due to lack of data) equation (1) was in addition estimated with $Emp_{i,t}$, an employment volume index available from OECD-Datastream, replacing $GDPR_{i,t}$ above.[18]

[15] More details about all other single-equation specifications we employed in estimating pass-through on import prices à la Campa-L.Goldberg (2002), including regression output and EViews programs, are available upon request.

[16] It is true that the Campa-L.Goldberg (2002) competitiveness proxy does not render itself to a self-evident interpretation. Without much details in their paper (p. 8, paragraph 2), these authors state that the variable they construct should capture the shifting relative price of a country's trading partners and use it as a consolidated export partners cost proxy. The benefit from this particular measure is that it is readily constructible from standard macrodata (such as IFS NEERs, REERs and CPIs). That is why, for comparability to their estimates of pass-through and given the lack of an easy substitute for it, we also use the Campa-L.Goldberg competitiveness proxy in our computations.

[17] We also used lags of real GDP in specification (1). However, this has not significantly affected the pass-through coefficients of interest in the present study, as they are reported in tables 4 and 5 in Appendix B and discussed in section 4.1. More detailed results are available upon request.

[18] Thus eliminating the problem of real GDP interpolation; yet introducing other problems, of course.

We also estimated all corresponding *quarterly*-based specifications (including an additional one with $Emp_{i,t}$ for Germany, since the OECD German employment volume index was available at this particular frequency as well), which differ from (1) in that the respective sums are $\sum_{k=0}^{4}$ for the two lagged explanatory variables and in that there is just one, but quarterly, AR term to correct for first-order serial correlation in the residuals.

Following the literature, in particular Campa and L.Goldberg (2002) and Choudhri, Faruquee and Hakura (2002), we focus in this paper on the *time profile* of pass-through. Pass-through is, consequently, defined by the *cumulative sum* of the coefficients to the $NEERInv_{i,t-k}$ variable up to a given lag k. In tables 4 and 5 in Appendix B we report, and in section 4.1 interpret, such pass-through on import prices – in effect, measuring *elasticities* given the log-difference functional form specified – within the horizon of 1 year.

To check for parameter stability, we next performed tests for structural changes. Looking, first of all, at the respective exchange rate graphs in Appendix A, we identified the most likely break points for each of the three countries. Thus, in the case of the inverse US NEER in Figure 1, two potential breakpoint candidates suggested from the data stood out. Until March 1985 the US dollar index trended down (appreciation), then – which should partly be related to the Plaza and Louvre accords – until June 1995 it trended up (depreciation), and finally – perhaps in anticipation of implementing the European Monetary Union (EMU) – the downward trend (appreciation) was restored. We therefore tested our US regression for break points in 1985:03 and 1995:06. The Chow *breakpoint* test could not reject the null of no structural break for any of these dates as well as for both of them taken together, no matter whether we used the F-statistic or the log likelihood ratio as test criteria. The Chow *forecast* test, in turn, produced somewhat less convincing results: it could not reject the null in 1995:06, no matter which of the two alternative test statistics we used; as to the null in 1985:03, it was definitely rejected by the log likelihood ratio (with an associated probability of 0.0000) but decisively not rejected by the F-statistic (with an associated probability of 0.9865). For Germany and Japan, the graphs of the inverse NEER in figures 2 and 3, respectively, show a coinciding (local) minimum (strongest currency) in April 1995; but both of the above-mentioned Chow tests could not reject the null of no structural break at that particular point in time. Given the rejection of structural changes at the most critical NEER-related – and, hence, pass-through relevant – points in our data set for Germany and Japan and the only partial and conflicting test results for the US case about a potential break in March 1985, we concluded that the Chow tests did not find any strong evidence for structural breaks in all three economies analyzed. We then tested for a breakpoint in each of the countries exactly at the split of our sample, i.e. in January 1991. As already said, the reason for a sample split at that particular point in time was to obtain equal (that is, with the *same* number of observations) and, hence, more comparable subsample periods, denoted "the 1980s" and "the 1990s". For the US and Germany, both the Chow *breakpoint* test and the Chow *forecast* test could not reject the null of no structural change in 1991:01 at all usual levels of significance (i.e. at 1%, 5% and 10%). For Japan, however, we obtained somewhat ambiguous results: more precisely, the log likelihood ratio statistic of the Chow breakpoint test rejected the null at 5% and 10% but not at 1%, whereas the alternative F-statistic test criterion could not reject the null at these conventional significance levels, yet rejected it at just above 11%; at the same time, the Chow forecast test rejected the null at all usual levels according to

First, related to how much the IPI is representative for aggregate economic activity. This point is particularly valid for the three countries in question, given the large services sector in them. Second, related to how much employment is responsive to short-run changes in the business cycle.

the log likelihood ratio but the F-statistic decisively could not reject the null (with an associated probability as high as 0.9681). Therefore, our sample split in January 1991, from which we report our pass-through measures further down, should not lead to any detrimental consequences with respect to parameter stability, in particular in the cases of the US and Germany. In the Japanese case, such a sample split appears, moreover, to coincide with a likely break in structure at the time point in question.

Single-equation OLS regressions like the one we began our analysis with are common in empirical research, as they provide at least a first, benchmark estimation. Moreover, OLS is often the estimator with the minimum variance. That is why, in addition to its simplicity, it has been applied in the earlier pass-through literature too. And for the same reason, as well as for comparability, we started by extracting measures of pass-through from a particular OLS specification, defended by its proponents as attempting to synthesize and build upon most previous studies. However, OLS is known to yield estimates which are biased, the more so in small samples, when a regressor is correlated with the error term. This situation seems quite likely for some of the right-hand side variables in (1). To deal with a *potential bias*, we next estimated the same equation by the usual alternative to OLS, namely two-stage least squares (TSLS), itself a special case of the instrumental variables (IV) method. We employed as "instruments" the same variables as in the Campa-L.Goldberg (2002) specification but all lagged once. Now our results changed more, and in no systematic pattern across countries or subsample periods. *En gros*, the tendencies we summarize as robust conclusions from our present work remained valid again for Germany and Japan, yet not for the US.[19] To address this issue and perform an extensive sensitivity analysis of our initial pass-through measures, we moved on to compare our OLS estimates with ones obtained from VARs, as described below.

3.2 Pass-Through Estimates from VAR Systems

Application of vector autoregressions (VARs) is another widely used method to estimate the dynamic effects of shocks. In measuring pass-through from VAR systems, we principally pursued two objectives. First, to base our work on the recent advances in the related literature, essentially building upon them. Second, to stick at the same time to a parsimonious representation, bearing in mind the intended and most efficient use of VAR modelling. We now "borrowed" our specification from another recent study which claims to avoid weaknesses of previous similar research, namely Choudhri, Faruquee and Hakura (2002), but modified their system to a "minimal" one for our purposes here and complemented their estimation method as we explain below.

Testing for Cointegrating Relations Before specifying our VARs, we first duly checked for possible cointegrating relations among the variables to enter our system pass-through estimation. There has been a lot of disagreement in the literature as to whether cointegrated VAR models should be specified or not, in general as well as particularly when measuring pass-through. Two problems Choudhri, Faruquee and Hakura (2002) relegate to respective footnotes concern unit root and cointegration tests. These authors assume all their series except the interest rate to be I(1) based on Augmented Dickey-Fuller (ADF) tests and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests. They also note to have tested for potential cointegration related to 5 theory-suggested interdependencies among their 7 endogenous variables (including purchasing power parity), which has not been found. Coricelli, Jazbec and Masten (2003) go to the other extreme in basing their pass-through estimates on the only recently studied I(2) cointegrated VAR model, claiming that most nominal price data tend to be integrated of order 2

[19] Further details, including the EViews programs, are available upon request.

(and identifying 3 cointegrating vectors in the 5-variable system common to all 4 countries in their sample). As we already noted, in the special case of transition economies such a statement is perhaps statistically well-grounded, yet generally it need not be true.

Taking all these considerations seriously into account, we tried to be explicit and consistent in performing and interpreting our unit root and cointegration tests. We first checked for stationarity of potential cointegrating relations suggested by theory,[20] such as the (logs of the) terms of trade (ToT), purchasing power parity (PPP), the quantity theory of money (QTM)[21] and the ratio of the import price index to the CPI. We were not able to reject unit roots in these relations using four different specifications of each of the ADF and Phillips-Perron (PP) tests, as documented in Table 6 in Appendix B.

Moreover, we supplemented this initial check by formal cointegration tests using Johansen's procedure. In particular, the *summary* test taking account of five possible specifications was applied. We generally found quite diverging results on the number of cointegrating vectors potentially linking the variables in our 4 theory-induced interdependencies referred to above as well as among the 5 time series we employ in our (nominal) VARs later, also duly selected given our objective to estimate pass-through at different price levels and the "constraint" for a parsimonious specification: import, export and consumer prices, the nominal exchange rate and narrow money (M1). The results from these tests are summarized in Table 7 in Appendix B.

Having no clear guidance on the number of possible cointegrating relations, we thus did not engage in attempting to set up reasonable cointegrated VAR models for our data. This has, moreover, ensured greater comparability between the respective estimates of pass-through via OLS and impulse responses from VARs we report in the present paper.

Orthogonalized VAR Impulse Responses The most straightforward way to run a VAR is if the researcher leaves it *unrestricted*. In fact, the only restriction in this case is the Cholesky ordering which predetermines impulse responses and variance decompositions. This is the approach in estimating pass-through preferred, for instance, by McCarthy (2000) and, in essence, Choudhri, Faruquee and Hakura (2002). In applying it to our choice of sample and variables, we first used pairwise Granger-causality tests[22] and prior intuition from economic theory to reduce the possible causal chains to a few most likely subsets of orderings. In a next step, we compared our orthogonalized impulse responses across the four specifications supported by the data, thus providing some sensitivity analysis of our VAR pass-through estimates. These turned out to be rather robust to the four orderings we identified from the Granger tests, which may be partly due to the generally low contemporaneous correlations between the variables in the system.[23] In addition, a generalized VAR estimation (to be commented later) finally confirmed that substantial errors related to our data-and-theory-informed selection of orderings would be unlikely.

The major benefit from using *unrestricted* VARs is that they remain (perhaps the only tool) *usable* when theoretical prescriptions for structural identification of the model are insufficient, if not contradictory or missing at all, as we believe is the case here. That is why we abstain in this paper from experimenting with *structural* VARs too.

The vector autoregressive (VAR) representation of the simultaneous equations model we apply can be compressed in the following general notation:

[20] As done in Choudhri, Faruquee and Hakura (2002).

[21] To be more precise, we tested a simplified version of it implying unitary velocity.

[22] Granger causality does not, however, provide information on *within*-month causality (I am grateful for this point to Hans Genberg). Nevertheless, it is the principal technique used in the VAR-related literature when it comes to determining the order of variables. Fortunately, our results proved not to be much sensitive to ordering.

[23] For the precise numbers, see Table 8 in Appendix B. A look into the table would also indicate a few exceptions in the pairwise correlations which are relatively high.

$$\underset{(n\times n)}{A(L)}\ \underset{(n\times 1)}{y_t} = \underset{(n\times 1)}{\varepsilon_t}, \tag{2}$$

where

$$A(L) \equiv A_0 - \sum_{k=1}^{\infty} A_k L^k$$

is a *one-sided* matrix polynomial. In (2), the exogenous shocks $\underset{(n\times 1)}{\varepsilon_t}$ are written as a distributed lag of current and lagged values of the endogenous variables $\underset{(n\times 1)}{y_t}$.

In our particular version of (2) $n = 5$, with the five variables making up the endogenous vector $\underset{(n\times 1)}{y_t}$ specified in four orderings (presented below), and the lag structure is approximated by a truncation at 12 ($k = 1, 2, ..., 12$) motivated by the monthly frequency of the data.

The corresponding vector moving average (VMA) representation of the system (2) from which our impulse response measures of pass-through are inferred after imposing Cholesky orthogonalization of $\underset{(n\times n)}{\Sigma} \equiv E(\varepsilon_t \varepsilon_t')$, the variance-covariance matrix of ε_t, is:

$$\underset{(n\times 1)}{y_t} = \underset{(n\times n)}{C(L)}\ \underset{(n\times 1)}{\varepsilon_t}, \tag{3}$$

with

$$C(L) \equiv \sum_{k=0}^{\infty} C_k L^k.$$

Hamilton (1994: chapters 11 and 12) and Watson (1994) provide perhaps the standard references on the above correspondence between VAR and VMA representations and the related impulse response and variance decomposition analysis.

As mentioned, like most VAR researchers we relied on pairwise Granger-causality tests to judge about the most likely ordering of the five variables involved in our unrestricted VAR specifications. The tests were performed for the raw data[24] as well as for the seasonally adjusted ones, when these latter enter instead the system regressions due to identified seasonality. The outcomes from the Granger tests are summarized in figures 7 (for the raw data) and 8 (with seasonal adjustment) in Appendix B. Looking into these figures, sort of country-specific yet to some extent generalizable, motivated us to concentrate on a (12-lag) VAR alternating the following four orderings of the five variables (in first log-differences with a constant included) for each of the three countries examined.

1. Money → exchange rate → import prices → export prices → inflation: this is the ordering most frequently suggested by the Granger-causality tests (see again figures 7 and 8 in Appendix B). In our notation:

$$M1_{i,t} \rightarrow NEERInv_{i,t} \rightarrow PMI_{i,t} \rightarrow PXI_{i,t} \rightarrow CPI_{i,t}.$$

2. Ordering is the same as in the specification above but with the exchange rate first and money second, as indicated by part of the Granger tests and in accordance with

[24] Because seasonal adjusment may have distorted the original relationship between the variables in the system and as a comparability check with respect to the same tests effected with the respective deseasonalized time series.

a popular central bank policy which pays some more attention (at least implicitly) to the exchange rate:

$$NEERInv_{i,t} \to M1_{i,t} \to PMI_{i,t} \to PXI_{i,t} \to CPI_{i,t}.$$

3. Essentially, we now impose theoretical priors on the ordering which was most supported by our data, i.e. the one reflected in the first specification. This is done by moving the CPI from last to first in the causal chain, under the logic that inflation is the primary, if not the only, objective of most contemporary central banks, notably including the three countries of our present pass-through study:

$$CPI_{i,t} \to M1_{i,t} \to NEERInv_{i,t} \to PMI_{i,t} \to PXI_{i,t}.$$

4. Ordering is the same as in the preceding specification but with the exchange rate coming before money, in conformity with certain circularity between the Granger-causality found for those variables (in particular, for Germany):

$$CPI_{i,t} \to NEERInv_{i,t} \to M1_{i,t} \to PMI_{i,t} \to PXI_{i,t}.$$

As noted earlier, our orthogonalized impulse response estimates of pass-through from the above four specifications have been relatively robust to ordering. This is reflected in the time profile up to the horizon of one year extracted from these impulse responses and summarized by the "interval" estimates (as defined by the lowest and the highest point estimates across our VAR orderings) in tables 9 through 14 in Appendix B, which we shall interpret in section 4. But before doing so, we would now emphasize another novel feature of our empirical strategy aimed at robustifying comparative pass-through measurement. It consists in also employing generalized VARs, the underlying theory for which is introduced next.

Generalized VAR Impulse Responses Building on Koop, Pesaran and Potter (1996), Pesaran and Shin (1998) proposed *generalized* impulse response analysis as an alternative to the traditional, *orthogonalized* one outlined above. The main virtue of generalized VAR modelling is that, unlike the traditional one, it does not require orthogonalization of shocks and is invariant to the ordering of variables. We finally benefited from this recent theoretical contribution to VAR analysis by applying it to our system pass-through estimates, as another check of robustness across methodology. As far as we know, pass-through has not yet been estimated using this particular approach.

For the sake of clarity, we here briefly summarize generalized VAR theory. For further details and formal proofs the interested reader may wish to look up in the original Pesaran and Shin (1998) paper.

Koop, Pesaran and Potter (1996) define the *generalized* impulse response function at horizon l of a vector like y_t we referred to above as:

$$GI_y\left(l, \delta, \Omega_{t-1}\right) = E\left(y_{t+l} \mid \varepsilon_t = \delta, \Omega_{t-1}\right) - E\left(y_{t+l} \mid \Omega_{t-1}\right), \tag{4}$$

where Ω_{t-1}, a non-decreasing information set, denotes the known history of the economy up to time $t-1$ and $\delta = (\delta_1, ..., \delta_m)'$ is some hypothetical $m \times 1$ vector of shocks hitting the economy at time t. Using (4) in (3) gives:

$$GI_y\left(l, \delta, \Omega_{t-1}\right) = C_l \delta,$$

which is independent of Ω_{t-1} but depends on the composition of shocks defined by δ.[25] Therefore the choice of hypothesized vector of shocks, δ, is central to the properties of the impulse response function. The traditional approach, suggested by Sims (1980), is to resolve this problem by surrounding the choice of δ via the Cholesky decomposition of $\Sigma = E\left(\varepsilon_t \varepsilon_t'\right)$, the variance-covariance matrix of ε_t:

$$PP' = \Sigma, \tag{5}$$

where P is an $m \times m$ lower triangular matrix. Then (3) can be rewritten as:

$$y_t = \sum_{k=0}^{\infty} \left(C_k P\right) \left(P^{-1} \varepsilon_{t-k}\right) = \sum_{k=0}^{\infty} \left(C_k P\right) \xi_{t-k}, \qquad t = 1, 2, ..., T,$$

such that $\xi_t = P^{-1} \varepsilon_t$ are orthogonalized, namely $E\left(\xi_t \xi_t'\right) = I_m$. Hence the $m \times 1$ vector of the orthogonalized impulse response function of a unit shock to the jth equation on y_{t+l} is given by:

$$\psi_j^o\left(l\right) = C_l P e_j, \qquad l = 0, 1, 2, ..., \tag{6}$$

where e_j is an $m \times 1$ selection vector with unity as its jth element and zeros elsewhere.

The alternative approach to that of Sims (1980) proposed by Pesaran and Shin (1998) consists in using (4) directly but, instead of shocking all the elements of the vector ε_t, to shock just one, say the jth, of its elements and integrate out the effects of other shocks using an assumed or the historical distribution of the errors. In this case one would have:

$$GI_y\left(l, \delta_j, \Omega_{t-1}\right) = E\left(y_{t+l} \mid \varepsilon_{jt} = \delta_j, \Omega_{t-1}\right) - E\left(y_{t+l} \mid \Omega_{t-1}\right).$$

Assuming further that ε_t has a multivariate normal distribution, Koop, Pesaran and Potter (1996) show that:

$$E\left(\varepsilon_t \mid \varepsilon_{jt} = \delta_j\right) = \left(\sigma_{1j}, \sigma_{2j}, ..., \sigma_{mj}\right)' \sigma_{jj}^{-1} \delta_j = \Sigma e_j \sigma_{jj}^{-1} \delta_j.$$

Therefore, the $m \times 1$ vector of the (unscaled) generalized impulse response of the effect of a shock in the jth equation at time t on y_{t+l} is:

$$\left(\frac{C_l \Sigma e_j}{\sqrt{\sigma_{jj}}}\right) \left(\frac{\delta_j}{\sqrt{\sigma_{jj}}}\right), \qquad l = 0, 1, 2... \,.$$

Finally, by setting $\delta_j = \sqrt{\sigma_{jj}}$, Pesaran and Shin (1998) derive the scaled generalized impulse response function:

$$\psi_j^g\left(l\right) = \sigma_{jj}^{-\frac{1}{2}} C_l \Sigma e_j, \qquad l = 0, 1, 2, ... \,. \tag{7}$$

This latter function measures the effect of one standard error shock to the jth equation at time t on expected values of the vector y at time $t + l$.

The generalized impulse response estimates of pass-through have coincided with our second orthogonalized VAR specification enumerated above, and are thus included in the range estimates reported in tables 9 through 14 in Appendix B. As shown by Pesaran and Shin (1998), such a coincidence can happen only when impulse responses are estimated for innovations in the first equation in the system, which is exactly the case of our second VAR specification. In all other cases, generalized and orthogonalized time profiles accounting for the system dynamics following a shock are theoretically different, with the generalized impulse response function robust to ordering but the orthogonalized one not.

[25] Pesaran and Shin (1998) note that this history invariance property of the impulse response is specific to linear systems and does not carry over to nonlinear ones.

4 Interpretation of Findings

We now discuss our estimates of NEER pass-through along three different stages in the pricing chain, i.e. on import prices, on export prices and on consumer prices, and in relation to their specificity across methodology, frequency, proxy, time and country.

4.1 Pass-Through on Import Prices

Single-Equation Methodology Comparing first our OLS findings about the empirical range of pass-through from the exchange rate to *import* prices in tables 4 and 5 in Appendix B, we are able to reveal the following main conclusions, along the several dimensions of our study highlighted below.

Across Frequency The OLS regression à la Campa-L.Goldberg (2002) we ran at different frequencies with the same underlying data[26] produced rather different pass-through estimates, mostly in terms of magnitudes at identical time horizons but also in terms of overall dynamic patterns. A general finding valid for Germany and Japan is that quarterly based estimates tend to somewhat *understate* pass-through relative to monthly based ones, especially over the very short run and for the whole sample and the 1980s subsample. This understatement, however, seems not very high, being of the order $10 - 20\%$ of the respective magnitudes, and is almost completely absorbed by the fourth quarter, thus resulting in converging estimates over one year. For the 1990s, Japanese quarterly and monthly estimates are really very close. As for the US, quarterly and monthly estimates differ, not too much for the whole sample and in the 1980s, but substantially in the 1990s (the quarterly magnitudes being 2 to 3 times *higher* than the corresponding monthly based ones).

Across Proxy There is also some difference in the time profile extracted, using OLS estimation, from the two *aggregate import price* proxies, *indexes* and *unit values* – mostly at the horizon of 2-3 months, hence 1 quarter; yet this difference likewise tends to diminish over longer horizons, 1 year in particular. Thus, for the cumulative pass-through on import prices, both our proxies result in quite close estimates, notably over the whole sample (109.0% using import price indexes and 110.3% using unit values of imports for Germany; and 100.0% and 104.3%, respectively, for Japan) and during the 1990s subperiod (57.0% and 57.3% for Germany; 52.8% and 53.2% for Japan). However, the slight *overstatement* of pass-through on import prices by OLS with unit values, almost imperceptible in the percentages we quoted, becomes more pronounced for the 1980s. To sum-up, the use of import unit values in place of price indexes seems to matter in terms of the *precise* magnitudes of NEER pass-through, especially in the short run of 1, 2 and 3 months (hence, 1 quarter), but not that much in capturing the *general* time profile.

As for *business cycle* controls in the Campa-L.Goldberg (2002) regression, using *industrial production* indexes or *employment* volume indexes instead of *real GDP* does not considerably affect our results either.[27] The interpolation of GDP-related data we used in our monthly pass-through measurement does not thus seem to matter much.

[26] In order not to overburden the paper with factual material, we have preferred to include in appendix only descriptive statistics and estimation results concerning our *monthly* series. The analogous information for the corresponding *quarterly* data as well as our EViews programs are, certainly, available upon request. Nevertheless, when discussing the sensitivity of our results across *frequency*, we also summarize the respective findings based on our quartely data, essentially comparing them to our corresponding monthly based conclusions.

[27] More details are available upon request.

Across Time Irrespective of the frequencies and proxies we employ, a common conclusion is that NEER pass-through on import prices has diminished sharply in the 1990s relative to the 1980s at almost all horizons up to one year, as documented in tables 4 and 5 in Appendix B. A notable exception to this general finding is just the pass-through on impact (i.e. in the first month – but not in the first quarter – following an exchange rate innovation) in the US, higher (but still quite low) in the 1990s (4.9%) than in the 1980s (2.5%). One of the principal reasons behind such a secular phenomenon could be a shift to a higher extent of CCP, or – which is similar – to an increased pricing-to-market behavior by monopolistic firms competing strategically in today's globalizing economy. As mentioned in the introduction, other papers of the late 1990s such as, for instance, Friberg and Vredin (1997) had already challenged Grassman's law derived from data in the 1960s and early 1970s by finding empirically a growing role for PTM, itself first predicted and theoretically justified by Krugman (1987).

Across Country The interesting but more or less known result from the cross-country comparison of our single-equation estimates of exchange rate pass-through on import prices is the *very low* pass-through – along all studied horizons – in the US relative to Germany and Japan. Only $\frac{1}{4}$ of a NEER change is estimated to be passed on to import prices over one year in the US and only about 4% in the first month, during the whole sample period as well as (a little bit more) in the 1990s subsample. By contrast, our estimates for Germany and Japan present evidence for a virtually *full* pass-through on import prices over the same horizon of one year within the total sample, with more than half of the cumulative change happening in the first month after the shock.

How Do Our Results Compare to Those in Campa-L.Goldberg (2002)? Our NEER pass-through elasticities on import prices obtained along the Campa-L.Goldberg (2002) OLS methodology but with *monthly* data and a corresponding specification, equation (1), are almost identical for the US, not much different for Japan and kind of exaggerated for Germany when compared to the *quarterly* based measures at the relevant horizons the mentioned authors report.[28] At one quarter Campa and L.Goldberg (2002) obtain[29] 18.4% for the US, 49.7% for Germany and 84.1% for Japan, within their whole sample of 1975-1999; at a horizon of one year the respective pass-through on import prices they find is 29.2%, 73.4% and 117.7%. Our own estimates are 18.3% and 24.4% for the US, 86.8% and 109.0% for Germany and 67.8% and 100.0% for Japan, using price indexes (as noted, employing unit values in this case would not change much).

VAR System Methodology To be able to directly compare our impulse response estimates of pass-through from the NEER to import prices obtained using VARs and documented in tables 9 and 10 with those obtained via OLS in tables 4 and 5 (all found in Appendix B), we applied a simple but informative transformation to the response values at all time horizons. This transformation consists in normalizing all impulse responses to an exchange rate innovation of one standard deviation by the magnitude of that same standard deviation. It results in a pass-through elasticity measured in percentage changes, just as in the case when we used first differences of natural logs to specify our OLS regressions. In effect, our VAR pass-through estimates quoted below continue to express what part (in %) is passed on to various price proxies following a unit change in the NEER, as it was until now. Moreover, with the help of this transformation

[28] Due to a lesser similarity/consistency of our OLS specification and sample with the one summarized in P.Goldberg and Knetter (1997), we would not engage here in comparing our quantitative findings with theirs.

[29] Cf. their Appendix Table 1, p. 29.

we can judge to what extent the econometric method applied (single-equation OLS vs. simultaneous VAR system, in particular) may affect our principal findings along the several dimensions of the present empirical analysis.

Across Frequency Turning back to the frequency dimension, we could sum up the following main conclusions from the VARs we ran. Estimates of (cumulative) pass-through at the same time horizon – e.g. one, two, three and four quarters – obtained from quarterly data are generally *lower* than the corresponding estimates based on monthly series. This is particularly true for the whole sample period and the 1980s subsample and for Germany and Japan. The US monthly vs. quarterly based estimates do not diverge a lot, for all subperiods and for all stages in the pricing chain.

An interesting observation which comes out from our monthly pass-through estimates – but impossible to be captured at a quarterly frequency – is related to a kind of short-run dynamics of price adjustment to exchange rate changes, rather common across stages in the pricing chain, variable proxies, time periods and country cases. There is a "dive" in our pass-through estimates, more frequently in the third month and less frequently in the second month. It usually comes after a "spike", generally in the first or second month. Such a pattern in the initial dynamics of pass-through obviously exhibits some "overshooting", which appears typical for the economies we focused on.

Across Proxy Except for Germany in the 1990s, the use of one or the other of our two proxies of aggregate prices of imports in the VARs did not appear to change much, as it was with our OLS estimates. More precisely, unit value inferred impulse response measures tend to slightly *understate* pass-through on import prices in the shorter run (up to one quarter). Sometimes this underestimation is complemented by a weak exaggeration in the longer run (one year).

Across Time For all countries and no matter the frequency or proxy, VAR-estimated pass-through on import prices has decreased in the 1990s relative to the 1980s – weakly for the US, dramatically for Germany and, to a lesser extent, Japan – and at all horizons (except in the very short run in the US). This conclusion generally accords with our OLS estimates. However, the magnitude of the empirical range of pass-through measured via OLS vs. VARs as well as, consequently, the extent of decrease in pass-through in the recent decade differ across methodologies. As a result, US estimates from OLS and VARs largely coincide across all sample periods and horizons. The same is true for Japan and Germany in the 1990s, but not in the 1980s and, therefore, over the whole sample.

Across Country In the US, the single-equation *point* estimate à la Campa-L.Goldberg (2002) is most of the time *inside* the system *range* estimate summarizing the four alternative orderings of our VARs. As to the generalized impulse response measures, they coincide with our orthogonalized impulse response findings when ordering with the exchange rate coming first (as in our second VAR specification) is effected, in compliance with the theoretical result by Pesaran and Shin (1998) mentioned earlier. The generalized impulse response magnitudes are thus also included within the intervals reported in tables 9 and 10 in Appendix B. If there are some differences to distinguish between OLS and VAR pass-through estimates on import prices in the case of the US, these would concern the cumulative response at the longer horizons (3 quarters and 1 year) and mostly the 1990s subperiod (when VAR-obtained values are somewhat higher). Otherwise, our OLS and VAR measures of pass-through on import prices are practically unanimous in the US case. As we said, this is not so for Japan and Germany right after the first month following an exchange rate depreciation has elapsed. In cumulative terms over the horizon of one year, Japanese VARs tend to *overestimate* pass-through on import prices relative to OLS by about $\frac{1}{3}$ during the whole sample period as well as over the 1980s; German VARs exaggerate pass-through on import prices with respect to our OLS estimates roughly twice over the same horizon. However, in the 1990s subperiod both Japanese and German VARs extract from the data ranges of pass-through on import prices largely similar to those obtained via OLS.

4.2 Pass-Through on Export Prices

Looking now at the pass-through from exchange rate changes to *export* prices in tables 11 and 12 in Appendix B, we could summarize our findings in the following manner.

Across Frequency With respect to pass-through on export price levels, the frequency dimension of our study does not easily lend itself to a simple generalization. On the one hand, Germany and Japan seem again more similar between themselves, with the US standing out as a special case. But the fact that quarterly estimates tend to *understate* pass-through relative to monthly ones remains valid for Japan in the whole sample and its two subperiods (with less divergence compared to what we observed concerning import prices) as well as for Germany in the whole sample and during the 1980s. For the US a similar conclusion is true for the 1980s only, not for the whole sample and the 1990s.

Across Proxy Although again preserving some very general trends, the estimates resulting from *unit values* now produce time profiles that are quite dissimilar to (in fact, much steeper than) the corresponding estimates obtained from *price indexes*. Moreover, in the Japanese case, unit value estimates are indicative of falling pass-through on the prices of exports in the 1990s relative to the 1980s, while price index-based measures reverse this conclusion. In the case of Germany, estimates based on indexes present evidence for a pass-through that diminishes considerably in the 1990s relative to the 1980s, especially over the one-year horizon, whereas estimates from unit values indicate only a modest reduction. Our proxy check, therefore, flashes a red light: measurement problems involved in unit values and price indexes may impair, as here, the robustness of similar pass-through estimates.

Across Time As to the general trend of *declining* pass-through across time, both discussed proxies confirm this conclusion only for Germany; the exact magnitude of this decline, however, differs, as we noted above. With respect to the US, pass-through on export prices has somewhat *increased* in the 1990s relative to the 1980s: from $11.7-15.9\%$ to $16.5-17.6\%$ over a horizon of one year. The same tendency, but at a much higher pass-through magnitude, is true for Japan if price indexes are used in the VARs but not unit values, as already commented.

Our empirical findings thus indicate a considerably declining pass-through on import prices accompanied with more or less stable pass-through on export prices in the US, Germany and Japan. Observe that a simple two-country model of the types used in traditional and new open-economy macroeconomics would not capture such a *pass-through asymmetry*. The reason is that two-country models impose symmetric imports and exports: what is exports for the first economy is, by necessity, imports for the second one. A trading system in the real world remains closed too, but is not restricted to two countries only, so asymmetries on a bilateral basis are not excluded (and are often a feature of the data). Nevertheless, an interpretation of this asymmetric pass-through on import and export prices we would propose is, at least in part, consistent with our theoretical work in Mihailov (2003 a, b). It boils down to the following trends in the price-setting behavior of monopolistically competitive producers and/or exporters: foreign exporters to the US, Germany and Japan have tried to maintain their shares in the huge markets of these economies throughout the 1990s by (i) more recourse to pricing-to-market, i.e. to exports priced according to CCP, and (ii) less pass-through from exchange-rate changes to the prices of that fraction of their exports which is denominated in the respective own national currency, i.e. priced according to PCP; at the same time, exporters from these three major economies to relatively smaller markets (of many other countries) have been more reluctant, when pricing exports, to shift from their domestic but world-wide accepted currency to foreign currencies (in particular, such that are of marginal significance in global forex markets).

Across Country The empirical range of pass-through across countries is, again, quite varied when pass-through on export prices is analyzed. The lowest pass-through is in the US (like it was with pass-through on import prices), of the order of $13.7 - 15.4\%$ at the one year horizon for the whole sample period. Japan exhibits the highest pass-through on export prices for that same horizon and period, $69.3 - 70.0\%$, and Germany comes close to Japan, with $54.4 - 57.7\%$. The three interval estimates just quoted were those obtained via price indexes (using unit values instead would produce kind of opposite ranges for Germany and Japan).

4.3 Pass-Through on Consumer Prices

We finally compare our findings about the empirical range of pass-through to *consumer* prices, reported in tables 13 and 14 in Appendix B. Here several conclusions that hold in common for the three countries considered seem to be shaping out.

Across Frequency As far as NEER pass-through on consumer price levels is concerned, frequency largely does not matter. A general finding is that at this final stage in the pricing chain, relevant for consumers' decision-making and, hence, for any microfounded macroeconomic outcomes, pass-through is low to negligible.

Across Proxy The proxy employed in our impulse response estimates of pass-through on consumer prices does not matter either. For the whole sample and the 1990s (but much less so for the 1980s) empirical ranges along all respective horizons are very close in value, thus producing a very similar time profile in Germany and in Japan.

Across Time In Germany and Japan, proxies accord as well on the tendency towards a decline in the exchange rate pass-through on consumer prices in the 1990s compared to the 1980s. As to the US, there is strong evidence that this particular pass-through has been negligible at all time horizons, over the whole sample period and within each of the two subperiods.

Across Country A major conclusion is thus that there is nowadays a *practically nil* pass-through from exchange rate movements to consumer prices in all three countries examined.

How Do Our Results Compare to Those in Choudhri, Faruquee and Hakura (2002)? Using orthogonalized impulse responses from a somewhat different sample period and VAR specification with seven endogenous and two exogenous variables over quarterly data, Choudhri, Faruquee and Hakura (2002) measure the exchange rate pass-through at various stages of the pricing chain for the six non-US G-7 countries. Are our findings at the relevant horizons similar to theirs?[30] Generally yes, mostly concerning consumer prices for both Japan and Germany and at both horizons of principal interest, one quarter and one year, as well as for Japan at all three levels in the pricing chain and at both mentioned time spans. The latter three authors report[31] pass-through on *import* prices of 80% at one quarter and 134% at one year for Japan and 39% and 77%, respectively, for Germany; our corresponding VAR interval estimates (employing price indexes) are $82.1 - 82.3\%$ and $137.8 - 141.2\%$ for Japan and $94.0 - 100.6\%$ and $205.0 - 219.6\%$ for Germany. Pass-through on *export* prices is, correspondingly, 50% and 50% for Japan and 3% and 16% for Germany in Choudhri *et al.* against our estimates of $74.6 - 74.7\%$ and $69.3 - 70.0\%$ for Japan and $21.3 - 22.2\%$ and $54.4 - 57.7\%$ for Germany. Finally, pass-through from exchange rate changes to *consumer* prices is measured by the three authors at -1% and 4% for Japan and 15% and 20% for Germany while our estimates are, respectively, $1.2 - 1.3\%$ and $6.0 - 6.2\%$ for Japan $1.4 - 4.2\%$ and $15.0 - 21.4\%$ for Germany.

5 Concluding Comments

The present paper built on some empirical implications of the theoretical NOEM literature as well as on studies of exchange rate pass-through using macrodata to measure and interpret the likely range of this phenomenon in three leading national economies in the world, namely the US, Germany and Japan. We obtained results employing various methods and specifications, and containing a number of interesting aspects to analyze. Focusing on monthly data to comply with a consensual span of predominant real-world price level stickiness, we inferred pass-though estimates that are broadly similar – when expressed in quarterly terms – to those extracted in earlier related papers, notably from OLS in Campa-L.Goldberg (2002) and from VARs in Choudhri, Faruquee and Hakura (2002). Yet the following novel features of our work, as well as some key differences along its several dimensions, are worth emphasizing.

An overall conclusion is that the empirical range of exchange rate pass-through on prices varies across (i) economies, (ii) data frequencies, (iii) periods of time, (iv) methods of estimation, (v) aggregate price measures, (vi) stages along the pricing chain and (vii) horizons of analysis. Any generalization thus needs to be careful. Yet abstracting from the specificity of some features of pass-through we commented in detail above, we would like to stress at least three important and rather robust results from our empirical analysis. First, in the three countries we examined, pass-through on import prices has considerably declined in the 1990s relative to the 1980s; but pass-through on export prices has, in essence, remained the same, although with certain country-specific nuances: more precisely, it has somewhat increased in the US, stayed flat in Japan and slightly decreased in Germany; as far as consumer prices are concerned, exchange rate pass-through seems to be nowadays practically negligible over all horizons of up to one year. Grassman's law evoked in the introductory part thus appears to be "weakening" by the end of the 20th century relative to the last decade of the Bretton-Woods era. Second, the econometric method and the measurement proxy used matter for the *precise* magnitudes and time patterns, yet they often – but not always – accord on the *general* trends. Third, the US is quite a particular economy, with import and, hence, consumer price levels that are amazingly insensitive to US dollar depreciations.

As far as our focus on the frequency dimension of pass-through estimates is concerned, a general insight from performing the same calculations with monthly as well as with (corresponding) quarterly data is that when passing from the higher to the

[30] Due to a lesser similarity/consistency of our VAR specification and sample with those in McCarthy (2000), we would not compare here our pass-through ranges with his related results.

[31] In their Table 1, p. 23.

lower frequency a lot of short-term dynamics is lost, partly due to an "averaging out" effect. When monthly fluctuations are strong, the difference in estimates from monthly vs. quarterly data should therefore be substantial. Conversely, for less volatile monthly data, quarterly estimates should offer good approximations. This intuitive logic is supported by the evidence for a difference in the magnitudes, and sometimes in the trends, of estimated pass-through at different data frequency the present empirical work revealed.

A Data: Definitions, Graphs, Descriptive Statistics

A.1 Definitions of the Data

Country Codes

- US: United States
- BD: Germany
- JP: Japan

Data Sources

- IFS: *International Financial Statistics*, International Monetary Fund (IMF), via Datastream
- MEI: *Main Economic Indicators*, Organization for Economic Cooperation and Development (OECD), via Datastream

Variable Codes and Sources

- PMI: import price index, IFS (USI76.X.F, BDI76.X.F, JPI76.X.F)
- PMU: unit value of imports, IFS (BDI75...F, JPI75...F)
- PXI: export price index, IFS (USI76...F, BDI76...F, JPI76...F)
- PXU: unit value of exports, IFS (BDI74...F, JPI74...F)
- NEER: nominal effective exchange rate index, IFS (USI..NEUE, BDI..NEUE, JPI..NEUE)
- REER: real effective exchange rate index, IFS (USI..REUE, BDI..REUE, JPI..REUE)
- NEERInv: inverse of the nominal effective exchange rate index $\equiv \frac{1}{\text{NEER}}$
- CPI: consumer price index, IFS (USI64...F, BDI64...F, JPI64...F)
- Nominal GDP (quarterly), IFS (USI99B.CB, BDI99B.CB, JPI99B.CB)
- GDP deflator (quarterly), IFS (USI99BIRH, BDI99BIRH, JPI99BIRH)
- Real GDP: nominal GDP divided by the GDP deflator $\equiv \frac{\text{nominal GDP}}{\text{GDP deflator}}$
- IPI: industrial production index, IFS (USI66..IG, BDI66..IG, JPI66..IG)
- Employment (monthly, for the US and Japan): employment volume index, MEI (USOEM040G, JPOEM040G)
- Employment (quarterly, for Germany): employment volume index, MEI (BDOEM040H)
- C-G Cost: Campa-L.Goldberg (2002) cost competitiveness proxy $\equiv \frac{\text{NEER} \times \text{CPI}}{\text{REER}}$
- M1 (for the US and Japan): narrow money M1, IFS (USI34...A, JPI34...A)
- CC (for Germany): currency in circulation, IFS (BDL34A.NA)
- DD (for Germany): demand deposits, IFS (BDL34B.NA)
- M1 (for Germany): narrow money $\equiv$ CC + DD

Notation for Transformed Data

- SA: seasonally adjusted (via the Census X12 procedure) series used in estimation, after finding evidence for seasonality
- dl: first difference in natural logarithms of a series (i.e. percentage change)

A.2 Graphs of the Data

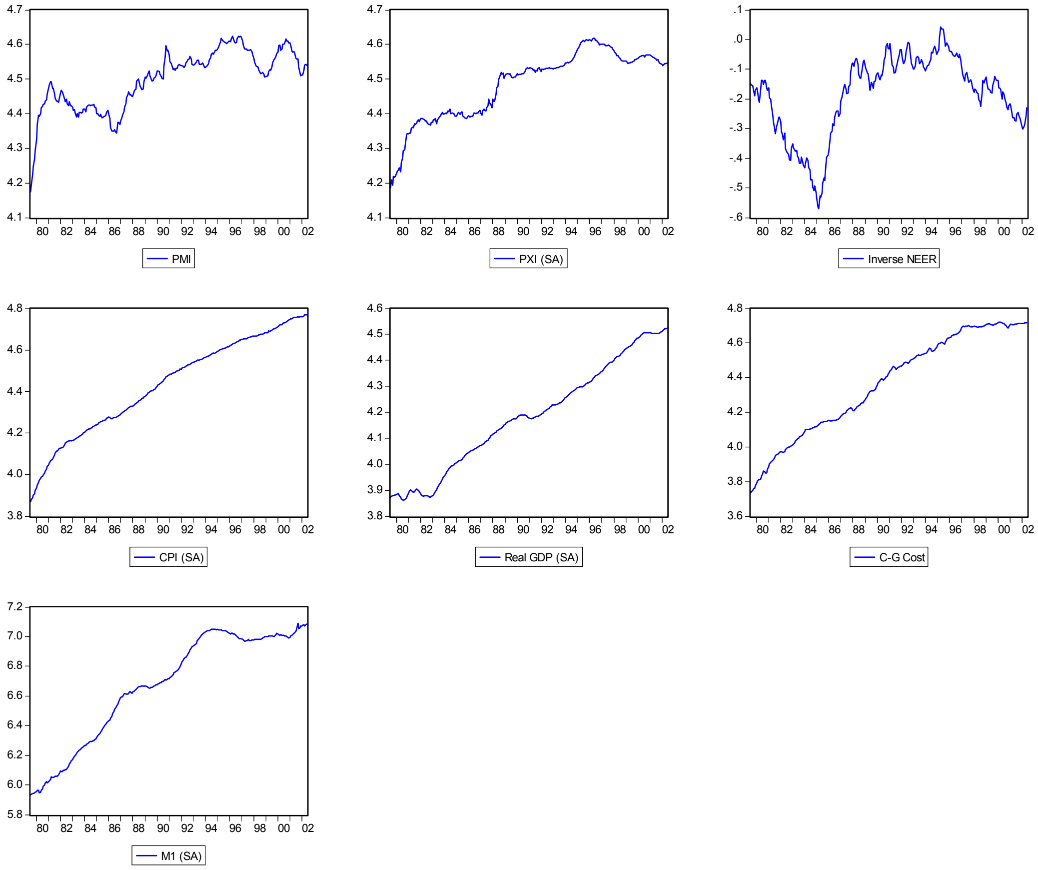

Figure 1: Graphs of (the Natural Logarithms of) the US Time Series Used in the Pass-Through Estimations: whole sample (1979:07-2002:06, 276 monthly observations)

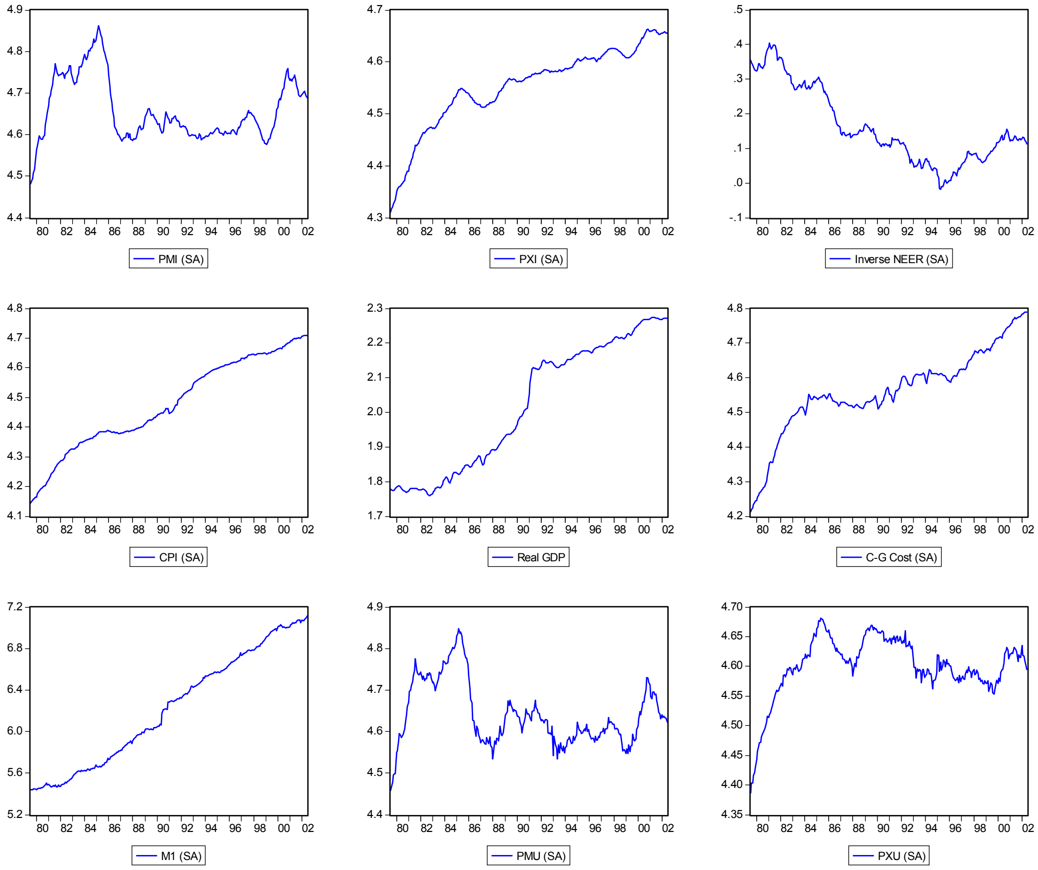

Figure 2: Graphs of (the Natural Logarithms of) the German Time Series Used in the Pass-Through Estimations: whole sample (1979:07-2002:06, 276 monthly observations)

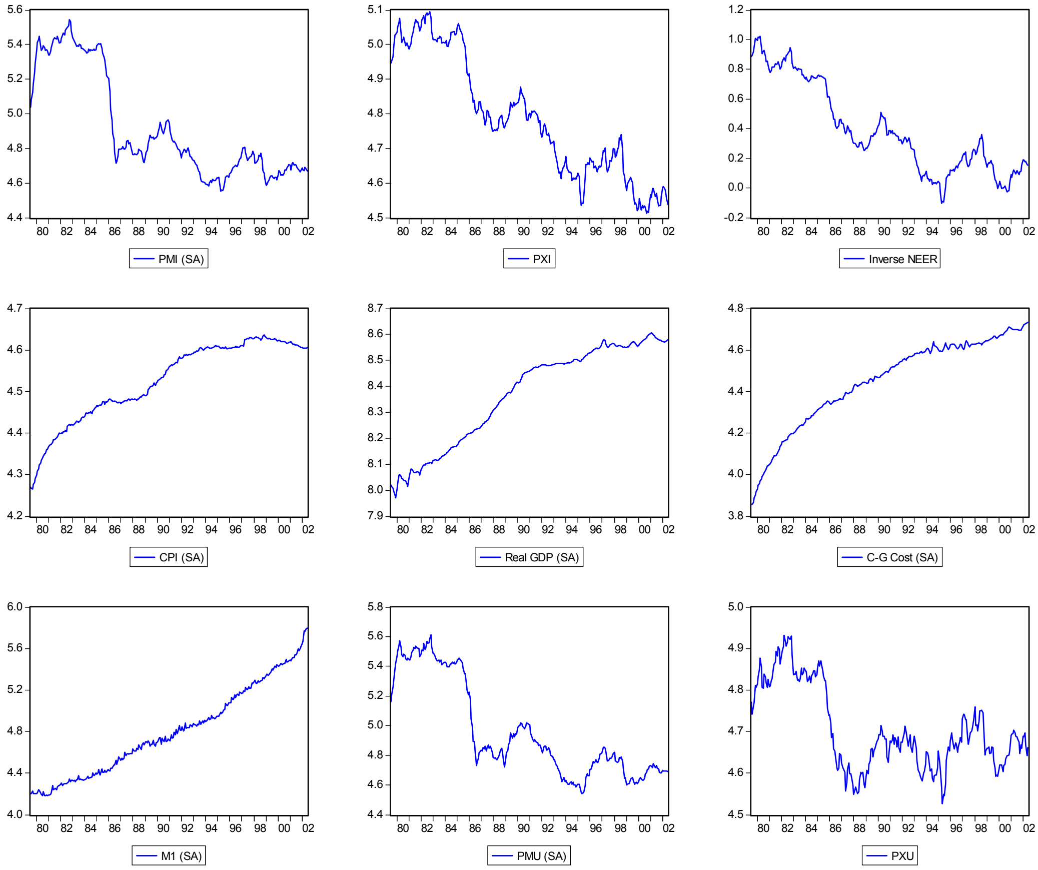

Figure 3: Graphs of (the Natural Logarithms of) the Japanese Time Series Used in the Pass-Through Estimations: whole sample (1979:07-2002:06, 276 monthly observations)

A.3 Descriptive Statistics of the Data

	PMI	PXI (SA)	Inverse NEER	CPI (SA)	Real GDP (SA)	C-G Cost (SA)	M1 (SA)
Whole Sample: 1979:07 - 2002:06							
Mean	0.0014	0.0013	-0.0002	0.0033	0.0024	0.0036	0.0042
Median	0.0010	0.0007	-0.0025	0.0028	0.0026	0.0032	0.0039
Maximum	0.0414	0.0292	0.0570	0.0140	0.0093	0.0193	0.0418
Minimum	-0.0265	-0.0164	-0.0498	-0.0049	-0.0073	-0.0099	-0.0340
Std. Dev.	0.0103	0.0062	0.0185	0.0028	0.0026	0.0051	0.0068
Skewness	0.7532	1.2333	0.2946	1.3028	-0.7841	0.3089	-0.0401
Kurtosis	4.7477	7.2955	3.0668	5.6308	4.8702	3.3070	8.9630
Jarque-Bera	61.2238	282.1515	4.0433	157.6730	68.5053	5.4738	408.9822
Probability	0.0000	0.0000	0.1324	0.0000	0.0000	0.0648	0.0000
Sum	0.3943	0.3650	-0.0604	0.9118	0.6524	0.9950	1.1603
Sum Sq. Dev.	0.0289	0.0106	0.0941	0.0022	0.0019	0.0072	0.0126
Observations	276	276	276	276	276	276	276
1980s Subsample: 1979:07 - 1990:12							
Mean	0.0032	0.0025	0.0010	0.0045	0.0022	0.0051	0.0057
Median	0.0017	0.0016	-0.0007	0.0037	0.0026	0.0044	0.0058
Maximum	0.0414	0.0292	0.0570	0.0140	0.0093	0.0193	0.0214
Minimum	-0.0265	-0.0164	-0.0498	-0.0049	-0.0073	-0.0088	-0.0164
Std. Dev.	0.0122	0.0080	0.0204	0.0033	0.0033	0.0052	0.0059
Skewness	0.7001	0.7545	0.2906	0.7200	-0.6920	0.3119	-0.3230
Kurtosis	3.6623	4.5980	2.6828	3.7839	3.7820	3.3685	3.8148
Jarque-Bera	13.7964	27.7752	2.5209	15.4567	14.5319	3.0181	6.2167
Probability	0.0010	0.0000	0.2835	0.0004	0.0007	0.2211	0.0447
Sum	0.4382	0.3502	0.1404	0.6207	0.3092	0.6984	0.7916
Sum Sq. Dev.	0.0205	0.0088	0.0568	0.0015	0.0015	0.0037	0.0047
Observations	138	138	138	138	138	138	138
1990s Subsample: 1991:01 - 2002:06							
Mean	-0.0003	0.0001	-0.0015	0.0021	0.0025	0.0021	0.0027
Median	0.0000	0.0000	-0.0031	0.0021	0.0026	0.0020	0.0020
Maximum	0.0194	0.0092	0.0437	0.0067	0.0058	0.0141	0.0418
Minimum	-0.0233	-0.0076	-0.0494	-0.0020	-0.0026	-0.0099	-0.0340
Std. Dev.	0.0074	0.0032	0.0164	0.0014	0.0018	0.0046	0.0073
Skewness	-0.3440	0.2148	0.1707	0.0779	-0.3136	0.1568	0.3255
Kurtosis	3.2934	3.1210	3.4153	3.7047	2.6167	2.9255	11.9605
Jarque-Bera	3.2163	1.1454	1.6620	2.9947	3.1064	0.5973	464.1032
Probability	0.2003	0.5640	0.4356	0.2237	0.2116	0.7418	0.0000
Sum	-0.0439	0.0148	-0.2008	0.2911	0.3432	0.2967	0.3687
Sum Sq. Dev.	0.0076	0.0014	0.0369	0.0003	0.0004	0.0029	0.0072
Observations	138	138	138	138	138	138	138

Figure 4: Descriptive Statistics of (the First Differences in Natural Logarithms of) the US Time Series Used in the Pass-Through Estimations

	PMI (SA)	PXI (SA)	Inv. NEER (SA)	CPI (SA)	Real GDP	C-G Cost (SA)	M1 (SA)	PMU (SA)	PXU (SA)
Whole Sample: 1979:07 - 2002:06									
Mean	0.0008	0.0013	-0.0009	0.0021	0.0018	0.0021	0.0061	0.0007	0.0008
Median	0.0010	0.0010	0.0000	0.0016	0.0013	0.0019	0.0055	0.0001	0.0011
Maximum	0.0325	0.0153	0.0201	0.0148	0.0330	0.0229	0.1101	0.0512	0.0309
Minimum	-0.0271	-0.0044	-0.0298	-0.0192	-0.0114	-0.0186	-0.0278	-0.0707	-0.0265
Std. Dev.	0.0089	0.0025	0.0075	0.0028	0.0044	0.0059	0.0124	0.0142	0.0082
Skewness	-0.0106	0.9954	-0.4155	-0.4188	2.7623	0.1917	2.7914	-0.1264	-0.1340
Kurtosis	3.6190	6.1576	3.8152	17.5472	19.6336	4.4596	23.1006	5.3884	4.1461
Jarque-Bera	4.4117	160.2334	15.5830	2441.7026	3532.7668	26.1884	5004.8099	66.3387	15.9314
Probability	0.1102	0.0000	0.0004	0.0000	0.0000	0.0000	0.0000	0.0000	0.0003
Sum	0.2247	0.3514	-0.2437	0.5717	0.4972	0.5842	1.6748	0.1812	0.2100
Sum Sq. Dev.	0.0218	0.0018	0.0153	0.0021	0.0052	0.0095	0.0420	0.0557	0.0185
Observations	276	276	276	276	276	276	276	276	276
1980s Subsample: 1979:07 - 1990:12									
Mean	0.0013	0.0019	-0.0018	0.0024	0.0018	0.0026	0.0062	0.0015	0.0019
Median	0.0014	0.0018	-0.0014	0.0018	0.0021	0.0027	0.0036	0.0020	0.0017
Maximum	0.0325	0.0153	0.0201	0.0118	0.0135	0.0229	0.1101	0.0512	0.0192
Minimum	-0.0271	-0.0044	-0.0259	-0.0026	-0.0114	-0.0138	-0.0278	-0.0406	-0.0155
Std. Dev.	0.0108	0.0029	0.0076	0.0025	0.0039	0.0062	0.0147	0.0142	0.0065
Skewness	-0.1074	0.8619	-0.2621	0.7983	-0.1409	0.2336	3.2712	0.1249	0.1632
Kurtosis	2.8952	5.3773	3.5594	4.2142	3.9325	4.2246	22.3350	3.6233	3.2053
Jarque-Bera	0.3285	49.5833	3.3794	23.1333	5.4559	9.8776	2395.7122	2.5930	0.8551
Probability	0.8485	0.0000	0.1846	0.0000	0.0654	0.0072	0.0000	0.2735	0.6521
Sum	0.1772	0.2656	-0.2495	0.3274	0.2493	0.3649	0.8492	0.2030	0.2582
Sum Sq. Dev.	0.0160	0.0012	0.0079	0.0009	0.0020	0.0053	0.0296	0.0274	0.0057
Observations	138	138	138	138	138	138	138	138	138
1990s Subsample: 1991:01 - 2002:06									
Mean	0.0003	0.0006	0.0000	0.0018	0.0018	0.0016	0.0060	-0.0002	-0.0003
Median	0.0002	0.0004	0.0009	0.0015	0.0011	0.0013	0.0062	-0.0008	-0.0005
Maximum	0.0182	0.0058	0.0169	0.0148	0.0330	0.0196	0.0390	0.0474	0.0309
Minimum	-0.0214	-0.0033	-0.0298	-0.0192	-0.0043	-0.0186	-0.0241	-0.0707	-0.0265
Std. Dev.	0.0065	0.0019	0.0072	0.0030	0.0048	0.0055	0.0095	0.0143	0.0095
Skewness	0.0528	0.1613	-0.5735	-1.0367	4.1784	0.0663	0.2422	-0.3665	-0.0348
Kurtosis	3.6373	2.7312	4.3032	23.0755	24.8301	4.6507	5.1679	6.9923	3.7546
Jarque-Bera	2.3998	1.0138	17.3294	2342.1161	3141.7227	15.7691	28.3744	94.7347	3.3023
Probability	0.3012	0.6023	0.0002	0.0000	0.0000	0.0004	0.0000	0.0000	0.1918
Sum	0.0475	0.0858	0.0057	0.2444	0.2479	0.2193	0.8256	-0.0218	-0.0481
Sum Sq. Dev.	0.0057	0.0005	0.0071	0.0012	0.0032	0.0042	0.0125	0.0281	0.0124
Observations	138	138	138	138	138	138	138	138	138

Figure 5: Descriptive Statistics of (the First Differences in Natural Logarithms of) the German Time Series Used in the Pass-Through Estimations

	PMI (SA)	PXI (SA)	Inv. NEER (SA)	CPI (SA)	Real GDP	C-G Cost (SA)	M1 (SA)	PMU (SA)	PXU (SA)
Whole Sample: 1979:07 - 2002:06									
Mean	-0.0011	-0.0015	-0.0026	0.0013	0.0020	0.0032	0.0058	-0.0015	-0.0003
Median	-0.0002	0.0000	-0.0004	0.0007	0.0018	0.0029	0.0042	0.0007	0.0008
Maximum	0.0736	0.0486	0.0685	0.0156	0.0327	0.0273	0.1021	0.0746	0.0563
Minimum	-0.1035	-0.0614	-0.0931	-0.0060	-0.0164	-0.0178	-0.0643	-0.1322	-0.0651
Std. Dev.	0.0236	0.0163	0.0255	0.0032	0.0055	0.0065	0.0237	0.0270	0.0192
Skewness	-0.5970	-0.3818	-0.4655	1.2045	1.7768	0.1695	0.2574	-0.8611	-0.2295
Kurtosis	5.1758	3.6047	3.7489	5.5234	12.8803	4.1259	3.9226	5.7197	3.3532
Jarque-Bera	70.8343	10.9093	16.4181	139.9590	1267.8560	15.8990	12.8360	119.1737	3.8568
Probability	0.0000	0.0043	0.0003	0.0000	0.0000	0.0004	0.0016	0.0000	0.1454
Sum	-0.3124	-0.4055	-0.7304	0.3474	0.5604	0.8893	1.6077	-0.4265	-0.0711
Sum Sq. Dev.	0.1531	0.0731	0.1784	0.0029	0.0084	0.0115	0.1540	0.2012	0.1016
Observations	276	276	276	276	276	276	276	276	276
1980s Subsample: 1979:07 - 1990:12									
Mean	-0.0001	-0.0011	-0.0036	0.0022	0.0032	0.0048	0.0039	-0.0008	-0.0006
Median	0.0017	0.0000	-0.0006	0.0018	0.0029	0.0042	0.0035	0.0014	0.0008
Maximum	0.0736	0.0347	0.0610	0.0132	0.0327	0.0215	0.0712	0.0746	0.0390
Minimum	-0.1035	-0.0458	-0.0776	-0.0060	-0.0164	-0.0078	-0.0643	-0.1322	-0.0488
Std. Dev.	0.0281	0.0150	0.0245	0.0036	0.0069	0.0057	0.0239	0.0325	0.0188
Skewness	-0.6564	-0.4681	-0.6364	0.6893	1.4970	0.3189	0.0028	-0.9231	-0.3067
Kurtosis	4.6810	3.2456	3.6892	3.7690	9.0100	2.6641	3.3645	4.9742	2.9196
Jarque-Bera	26.1573	5.3869	12.0468	14.3276	259.2307	2.9875	0.7641	42.0083	2.2010
Probability	0.0000	0.0676	0.0024	0.0008	0.0000	0.2245	0.6825	0.0000	0.3327
Sum	-0.0170	-0.1466	-0.4959	0.2970	0.4366	0.6560	0.5404	-0.1079	-0.0854
Sum Sq. Dev.	0.1078	0.0309	0.0822	0.0018	0.0066	0.0044	0.0780	0.1445	0.0485
Observations	138	138	138	138	138	138	138	138	138
1990s Subsample: 1991:01 - 2002:06									
Mean	-0.0021	-0.0019	-0.0017	0.0004	0.0009	0.0017	0.0077	-0.0023	0.0001
Median	-0.0016	-0.0010	-0.0003	0.0001	0.0012	0.0017	0.0066	0.0003	0.0004
Maximum	0.0392	0.0486	0.0685	0.0156	0.0115	0.0273	0.1021	0.0543	0.0563
Minimum	-0.0585	-0.0614	-0.0931	-0.0048	-0.0119	-0.0178	-0.0518	-0.0660	-0.0651
Std. Dev.	0.0181	0.0175	0.0265	0.0025	0.0033	0.0069	0.0234	0.0203	0.0197
Skewness	-0.4487	-0.3040	-0.3450	1.9763	-0.6099	0.3199	0.5424	-0.4912	-0.1666
Kurtosis	3.4701	3.6712	3.7354	11.8419	5.5588	4.9245	4.3721	4.1799	3.6959
Jarque-Bera	5.9005	4.7166	5.8469	539.3653	46.2045	23.6492	17.5911	13.5551	3.4229
Probability	0.0523	0.0946	0.0537	0.0000	0.0000	0.0000	0.0002	0.0011	0.1806
Sum	-0.2954	-0.2589	-0.2346	0.0503	0.1238	0.2332	1.0672	-0.3185	0.0143
Sum Sq. Dev.	0.0450	0.0421	0.0959	0.0009	0.0015	0.0065	0.0750	0.0565	0.0531
Observations	138	138	138	138	138	138	138	138	138

Figure 6: Descriptive Statistics of (the First Differences in Natural Logarithms of) the Japanese Time Series Used in the Pass-Through Estimations

B Test and Estimation Results

	Test A	Test B	Test C	Test D	Conclusion
United States					
PMI	0	0	1	0	0
PXI	1	1	1	0	1
NEER	0	0	1	0	0
CPI	1	1	1	1	1
Real GDP	1	1	1	0	1
IPI	0	0	1	0	0
Employment	0	0	1	0	0
C-G Cost	1	1	1	1	1
M1	1	1	1	1	1
Germany					
PMI	1	1	1	0	1
PMU	1	1	1	0	1
PXI	1	1	1	0	1
PXU	1	1	1	1	1
NEER	1	1	1	0	1
CPI	1	1	1	1	1
Real GDP	0	0	1	0	0
IPI	0	0	1	0	0
C-G Cost	1	1	1	0	1
M1	1	1	1	1	1
Japan					
PMI	1	1	1	0	1
PMU	1	1	1	0	1
PXI	0	0	1	0	0
PXU	0	0	1	0	0
NEER	0	0	1	0	0
CPI	1	1	1	1	1
Real GDP	1	1	1	0	1
IPI	0	0	0	0	0
Employment	0	0	1	0	0
C-G Cost	1	1	1	0	1
M1	1	1	1	1	1

Table 1: Seasonality Test (Census X12) Results

Explanatory Note to Table 1: A: test for the presence of seasonality (coded 1 in the table when found) assuming stability; B: nonparametric test for the presence of seasonality assuming stability; C: moving seasonality test; D: combined test for the presence of identifiable seasonality.

	ADF	PP	KPSS	Conclusion
United States				
PMI	I(1)	I(1)	I(1)	I(1)
PXI (SA)	I(1):1,3; I(0):2,4	I(1)	I(1):1,3; I(2):2,4	I(1)
Inv. NEER	I(1)	I(1)	I(1):2,4; I(0):1,3	I(1)
CPI (SA)	I(1):1,3; I(2):2,4	I(1):1,3; I(0):2,4	I(1):1,3; I(2):2,4	I(1)
Real GDP (SA)	I(1)	I(1)	I(1)	I(1)
IPI	I(1)	I(1)	I(1)	I(1)
Employment	I(1)	I(1)	I(1):3,4; I(?):1,2	I(1)
C-G Cost (SA)	I(2)	I(1):1,3; I(0):2,4	I(1):1,3; I(2):2,4	I(1)
M1 (SA)	I(1)	I(1)	I(1):1,3; I(2):2,4	I(1)
Germany				
PMI (SA)	I(1)	I(1)	I(1):3,4; I(2):1,2	I(1)
PMU (SA)	I(1)	I(1)	I(1):3,4; I(2):1,2	I(1)
PXI (SA)	I(1)	I(1)	I(1):4; I(2):1,2,3	I(1)
PXU (SA)	I(1)	I(1)	I(2):3,4; I(?):1,2	I(1)
Inv. NEER (SA)	I(1)	I(1)	I(1):4; I(2):1,2,3	I(1)
CPI (SA)	I(1)	I(1):1,2; I(2):3,4	I(2):3,4; I(?):1,2	I(1)
Real GDP	I(1)	I(1)	I(1):3,4; I(?):1,2	I(1)
IPI	I(1)	I(1):1,2,3; I(0):4	I(1):3,4; I(?):1,2	I(1)
C-G Cost (SA)	I(1):1,2; I(0):3; I(2):4	I(1)	I(1):4; I(?):1,2,3	I(1)
M1 (SA)	I(1)	I(1)	I(1):1,2; I(?):3,4	I(1)
Japan				
PMI (SA)	I(1)	I(1)	I(1):3,4; I(2):1,2	I(1)
PMU (SA)	I(1)	I(1)	I(1):3,4; I(2):2; I(?):1	I(1)
PXI	I(1)	I(1)	I(1):3,4; I(2):1,2	I(1)
PXU	I(1)	I(1)	I(1)	I(1)
Inv. NEER	I(1)	I(1)	I(1):2,3,4; I(2):1	I(1)
CPI (SA)	I(1)	I(1)	I(2):3,4; I(?):1,2	I(1)
Real GDP (SA)	I(1):1,3; I(0):2,4	I(1)	I(1):3; I(2):4; I(?):1,2	I(1)
IPI	I(1):1,3; I(0):2,4	I(1):1,3; I(0):2,4	I(1):3; I(2):4; I(?):1,2	I(1)
Employment	I(1)	I(1):1,3; I(0):2,4	I(2):3,4; I(?):1,2	I(1)
C-G Cost (SA)	I(0)	I(1):1,2; I(2):3,4	I(1):1; I(2):3,4; I(?):2	I(?)
M1 (SA)	I(1):1,2,4; I(2):3	I(1):1,2,3; I(0):for 4	I(2):3,4; I(?):1,2	I(1)

Table 2: Stationarity Test Results

Explanatory Note to Table 2: For the *Augmented Dickey-Fuller (ADF)* tests, the most common (auto)*regression*-based method of testing for unit roots, specification 1 imposes constant, trend and 12 lags; 2 – constant and 12 lags; 3 – constant, trend and automatic selection of the lag structure using the modified Akaike criterion; 4 – constant and automatic selection of the lag structure using the modified Akaike criterion. For the *Phillips-Perron (PP)* and the *Kwiatkowski-Phillips-Schmidt-Shin (KPSS)* tests, which are among the most frequently used *nonparametric* (kernel) methods of testing for (non)stationarity, specification 1 imposes constant, trend and the AR spectral - GLS detrended data method of estimating the frequency zero spectrum; 2 – constant and the AR spectral - GLS detrended data method; 3 – constant, trend and the Bartlett kernel method of estimating the frequency zero spectrum; 4 – constant and the Bartlett kernel method.

(log-levels)	US (1969:01-)	Germany (1958:01-)	Japan (1957:01-)
Obstfeld-Rogoff (2000) *quarterly* sample (1982-1998)	0.31	0.43	0.29
our largest *quarterly* sample (-2003:1)	0.41	0.45	−0.62
our whole *quarterly* sample (1979:3-2002:2)	−0.08	0.95	0.81
our 1980s *quarterly* subsample (1979:3-1990:4)	0.10	0.90	0.89
our 1990s *quarterly* subsample (1991:1-2002:2)	0.22	0.76	−0.05
our largest *monthly* sample (-2003:03)	0.41	0.45	−0.62
our whole *monthly* sample (1979:07-2002:06)	−0.07	0.95	0.81
our 1980s *monthly* subsample (1979:07-1990:12)	0.10	0.90	0.88
our 1990s *monthly* subsample (1991:01-2002:06)	0.23	0.76	−0.06

Table 3: ToT-NEER Correlations

Pass-Through on the *Import Price Index* Following NEER Depreciation, %			
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)			
	United States	Germany	Japan
month 1	3.6	58.6	57.9
month 2	9.1	19.1	8.2
month 3	5.6	9.1	1.6
quarter 1	18.3	86.8	67.8
end-quarter 2, cumulative	19.9	97.9	86.9
end-quarter 3, cumulative	25.5	109.8	93.4
year 1, cumulative	24.4	109.0	100.0
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)			
	United States	Germany	Japan
month 1	2.5	71.9	64.0
month 2	9.5	16.7	17.2
month 3	9.0	15.1	7.4
quarter 1	21.0	103.7	88.6
end-quarter 2, cumulative	31.6	115.7	112.3
end-quarter 3, cumulative	39.6	127.3	113.3
end-year 1, cumulative	33.3	130.0	121.9
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)			
	United States	Germany	Japan
month 1	4.9	35.6	49.3
month 2	11.3	11.7	0.2
month 3	−1.5	−3.1	−5.2
quarter 1	14.6	44.1	44.3
end-quarter 2, cumulative	12.9	49.0	55.5
end-quarter 3, cumulative	15.4	57.3	55.2
end-year 1, cumulative	27.4	57.0	52.8

Table 4: OLS Estimates of the Pass-Through on Import Prices Obtained Using Import Price Indices

Pass-Through on the *Unit Value of Imports* Following NEER Depreciation, %		
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)		
	Germany	Japan
month 1	41.2	51.0
month 2	43.5	47.5
month 3	−4.0	−11.5
quarter 1	80.6	87.0
end-quarter 2, cumulative	116.6	97.4
end-quarter 3, cumulative	121.8	102.6
year 1, cumulative	110.3	104.3
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)		
	Germany	Japan
month 1	74.8	61.8
month 2	19.4	59.5
month 3	15.0	−17.8
quarter 1	109.2	103.4
end-quarter 2, cumulative	131.1	127.9
end-quarter 3, cumulative	165.1	115.6
end-year 1, cumulative	155.4	124.7
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)		
	Germany	Japan
month 1	−3.9	37.9
month 2	44.1	39.1
month 3	−35.2	−10.2
quarter 1	5.0	66.7
end-quarter 2, cumulative	54.5	58.3
end-quarter 3, cumulative	52.2	62.5
end-year 1, cumulative	57.3	53.2

Table 5: OLS Estimates of the Pass-Through on Import Prices Obtained Using Import Unit Values

(log-levels, largest sample)	United States	Germany	Japan
ToT	ns: ADF, PP	ns: ADF, PP	ns: ADF, PP
PPP	ns: ADF (except 2 at 5%), PP	ns: ADF, PP	ns: ADF, PP
QTM	ns: ADF, PP	ns: ADF, PP	ns: ADF, PP
Import Prices / CPI	ns: ADF, PP	ns: ADF, PP	ns: ADF, PP

Table 6: Cointegrating Relations Checks via Unit Root Tests

Explanatory Note to Table 6: ns means nonstationarity found by all four specifications (see the explanatory note to Table 2) of the ADF and the PP tests. To obtain a more direct relevance of results, the German data for the quantity theory of money (QTM) test as well as for the VAR tests in the whole sample and during the 1990s end in 1998:12, when the *IFS* DEM-denominated series for currency in circulation and demand deposits comprising M1 were discontinued.

(log-levels)	United States	Germany	Japan
ToT: largest sample	0, 1 or 2	0, 1 or 2	0
PPP: largest sample	0 (or 1)	0 or 1	0, 1 or 3
QTM: largest sample	0 or 1	0 or 1	0, 1 or 2
Import and Consumer Prices: largest sample	0, 1 or 2	0 or 1	0 or 1
VAR: whole sample (1979:07-2002:06; 276)	2, 4 or 5	2, 3 or 5	2, 3 or 5
same but *unit values* instead of price indexes	n.a.	1 (or 2)	2, 3, 4 or 5
VAR: 1980s subsample (1979:07-1990:12; 138)	0, 1 or 2	1 (or 2)	1, 2 or 5
same but *unit values* instead of price indexes	n.a.	1, 2, 3, 4 or 5	1, 4 or 4
VAR: 1990s subsample (1991:01-2002:06; 138)	1, 2 or 3	2	2 or 3
same but *unit values* instead of price indexes	n.a.	2	2, 3 or 4

Table 7: Cointegrating Relations Test Results from Johansen's Procedure

Explanatory Note to Table 7: The respective cells indicate the number of cointegrating relations identified by the five specifications of Johansen's procedure *summary* test in EViews; any number in parentheses means that it has been found just once. To obtain a more direct relevance of results, the German data for the quantity theory of money (QTM) test as well as for the VAR tests in the whole sample and during the 1990s end in 1998:12, when the *IFS* DEM-denominated M1 series was discontinued.

Whole Sample Period (July 1979 - June 2002, 276 observations)				
US				
	dlM1(SA)	dlNEERInv	dlPMI	dlPXI(SA)
dlNEERInv	0.12			
dlPMI	0.02	0.12		
dlPXI(SA)	0.13	−0.07	0.24	
dlCPI(SA)	−0.01	−0.03	0.50	0.26
Germany				
	dlM1(SA)	dlNEERInv(SA)	dlPMI(SA)	dlPXI(SA)
dlNEERInv(SA)	0.11			
dlPMI(SA)	−0.09	0.57		
dlPXI(SA)	−0.14	0.46	0.78	
dlCPI(SA)	−0.08	0.02	0.37	0.28
Japan				
	dlM1(SA)	dlNEERInv	dlPMI(SA)	dlPXI
dlNEERInv	0.04			
dlPMI(SA)	0.01	0.68		
dlPXI	0.05	0.93	0.68	
dlCPI(SA)	−0.13	−0.02	0.13	0.02

Table 8: Pairwise Monthly Correlation Matrix for the Estimated VARs

Pass-Through on the *Import Price Index* Following NEER Depreciation, %			
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)			
	United States	Germany	Japan
month 1	2.5 – 3.5	52.0 – 54.1	53.5 – 53.6
month 2	10.2 – 11.2	31.8 – 33.6	26.8 – 27.0
month 3	6.2 – 6.9	9.8 – 13.2	16.8 – 17.9
quarter 1	19.0 – 21.6	94.0 – 100.6	82.1 – 82.3
end-quarter 2, cumulative	15.8 – 19.4	121.7 – 132.8	107.5 – 108.3
end-quarter 3, cumulative	19.8 – 23.7	173.6 – 184.1	112.9 – 115.7
year 1, cumulative	21.6 – 26.7	205.0 – 219.6	137.8 – 141.2
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)			
	United States	Germany	Japan
month 1	1.2 – 2.6	65.7 – 72.0	53.9 – 54.5
month 2	10.2 – 11.8	35.2 – 38.7	32.8 – 33.4
month 3	8.3 – 9.5	18.1 – 24.3	0.0 – 0.6
quarter 1	19.9 – 23.6	120.3 – 133.7	87.5 – 87.8
end-quarter 2, cumulative	23.3 – 27.6	169.7 – 191.1	126.6 – 129.7
end-quarter 3, cumulative	32.1 – 41.3	240.7 – 260.9	149.0 – 149.3
end-year 1, cumulative	22.7 – 35.0	271.2 – 310.7	180.8 – 182.1
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)			
	United States	Germany	Japan
month 1	4.4 – 5.8	33.3 – 36.2	45.7 – 46.0
month 2	10.4 – 11.2	16.3 – 21.1	4.9 – 6.2
month 3	(−2.0) – (−1.5)	(−2.5) – (−1.3)	(−12.5) – (−11.2)
quarter 1	13.3 – 15.2	47.4 – 55.0	38.2 – 41.0
end-quarter 2, cumulative	7.5 – 10.5	31.3 – 40.4	37.6 – 38.8
end-quarter 3, cumulative	19.3 – 21.1	53.5 – 61.5	47.2 – 53.1
end-year 1, cumulative	31.8 – 33.3	58.1 – 71.8	54.5 – 60.1

Table 9: VAR Estimates of the Pass-Through on Import Prices Obtained Using Import and Export Price Indices

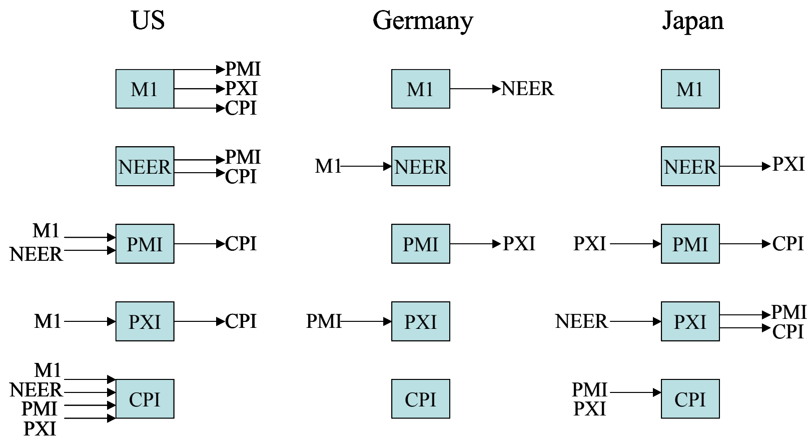

Figure 7: Summary of Pairwise Granger Causality Test Results (with 12 lags and at a 10% significance level threshold) for the Time Series Used in the VAR Pass-Through Estimates: *raw* data, largest sample (ending in 1998:12 for German pairs involving M1 – see the last sentence in the explanatory note to Table 7)

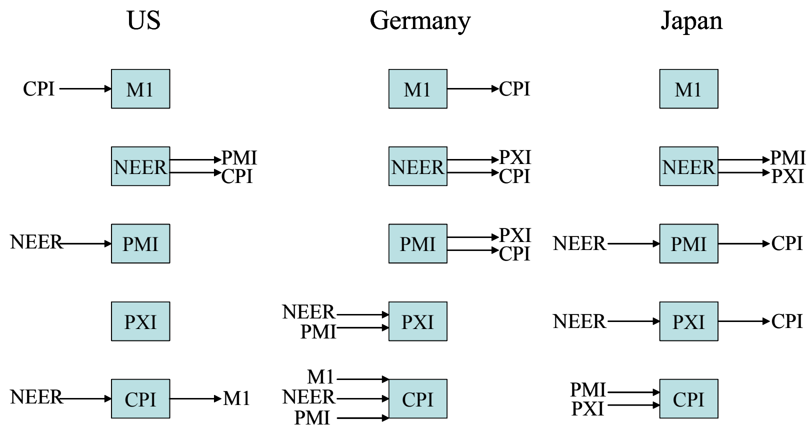

Figure 8: Summary of Pairwise Granger Causality Test Results (with 12 lags and at a 10% significance level threshold) for the Time Series Used in the VAR Pass-Through Estimates: *seasonally adjusted* data when seasonality found, largest sample (ending in 1998:12 for German pairs involving M1 – see the last sentence in the explanatory note to Table 7)

Pass-Through on the *Unit Value of Imports* Following NEER Depreciation, %		
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)		
	Germany	Japan
month 1	33.3 – 33.9	49.5 – 49.8
month 2	50.2 – 52.8	64.5 – 65.0
month 3	5.1 – 9.4	4.7 – 5.2
quarter 1	92.8 – 97.4	119.3 – 119.5
end-quarter 2, cumulative	140.2 – 146.7	120.5 – 122.1
end-quarter 3, cumulative	178.1 – 186.2	116.0 – 119.2
year 1, cumulative	205.9 – 221.7	130.6 – 134.7
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)		
	Germany	Japan
month 1	58.5 – 67.0	58.6 – 59.5
month 2	35.2 – 48.5	80.1 – 84.1
month 3	18.1 – 32.8	(−1.7) – (−0.7)
quarter 1	120.3 – 143.4	137.9 – 141.9
end-quarter 2, cumulative	169.8 – 187.2	173.0 – 178.0
end-quarter 3, cumulative	240.7 – 258.5	185.6 – 190.8
end-year 1, cumulative	268.6 – 314.8	214.3 – 220.4
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)		
	Germany	Japan
month 1	(−0.6) – 6.8	35.9 – 36.5
month 2	64.1 – 72.3	40.2 – 40.4
month 3	(−30.0) – 23.5	(−10.4) – (−8.3)
quarter 1	40.9 – 48.1	65.7 – 68.3
end-quarter 2, cumulative	107.3 – 110.0	39.1 – 41.7
end-quarter 3, cumulative	113.7 – 116.6	54.2 – 60.8
end-year 1, cumulative	148.8 – 152.9	60.0 – 67.4

Table 10: VAR Estimates of the Pass-Through on Import Prices Obtained Using Import and Export Unit Values

Pass-Through on the *Export Price Index* Following NEER Depreciation, %			
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)			
	United States	Germany	Japan
month 1	$(-2.8) - (-2.5)$	$12.7 - 13.2$	$55.0 - 55.3$
month 2	$1.3 - 1.5$	$6.6 - 7.0$	$21.4 - 21.6$
month 3	$4.4 - 4.6$	$2.0 - 2.2$	$(-20.0) - (-19.6)$
quarter 1	$3.0 - 3.5$	$21.3 - 22.2$	$74.6 - 74.7$
end-quarter 2, cumulative	$4.7 - 5.3$	$22.1 - 29.1$	$71.6 - 72.5$
end-quarter 3, cumulative	$13.3 - 14.4$	$42.7 - 45.3$	$60.5 - 62.0$
year 1, cumulative	$13.7 - 15.4$	$54.4 - 57.7$	$69.3 - 70.0$
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)			
	United States	Germany	Japan
month 1	$(-7.4) - (-6.0)$	$12.5 - 13.7$	$52.1 - 53.3$
month 2	$0.4 - 2.3$	$6.3 - 6.8$	$22.1 - 22.4$
month 3	$3.2 - 4.1$	$3.5 - 4.4$	$(-9.5) - (-9.3)$
quarter 1	$(-2.8) - (-1.8)$	$22.9 - 24.1$	$64.6 - 66.4$
end-quarter 2, cumulative	$(-2.4) - 0.3$	$37.5 - 39.0$	$64.1 - 65.0$
end-quarter 3, cumulative	$9.2 - 11.3$	$58.9 - 61.7$	$55.0 - 56.9$
end-year 1, cumulative	$11.7 - 15.9$	$75.1 - 78.6$	$58.8 - 59.4$
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)			
	United States	Germany	Japan
month 1	$2.0 - 2.5$	$11.5 - 11.9$	$50.7 - 51.3$
month 2	$2.3 - 2.4$	$5.8 - 7.1$	$10.5 - 12.1$
month 3	$2.2 - 2.4$	$(-0.9) - 0.1$	$(-3.9) - (-3.3)$
quarter 1	$6.7 - 7.0$	$16.4 - 19.0$	$58.0 - 59.6$
end-quarter 2, cumulative	$7.8 - 8.1$	$8.5 - 13.0$	$59.6 - 60.8$
end-quarter 3, cumulative	$13.2 - 14.2$	$11.0 - 15.7$	$69.2 - 73.9$
end-year 1, cumulative	$16.5 - 17.6$	$15.3 - 21.8$	$69.4 - 71.2$

Table 11: VAR Estimates of the Pass-Through on Export Prices Obtained Using Import and Export Price Indices

Pass-Through on the *Unit Value of Exports* Following NEER Depreciation, %		
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)		
	Germany	Japan
month 1	10.5 – 11.5	24.2 – 24.4
month 2	11.9 – 12.7	51.5 – 51.6
month 3	6.9 – 9.0	2.6 – 2.7
quarter 1	30.4 – 32.3	78.4 – 78.5
end-quarter 2, cumulative	40.8 – 44.0	61.6 – 62.6
end-quarter 3, cumulative	66.7 – 68.3	50.0 – 51.2
year 1, cumulative	75.7 – 78.1	57.2 – 58.8
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)		
	Germany	Japan
month 1	4.7 – 6.3	29.6 – 31.1
month 2	11.9 – 15.4	52.1 – 53.1
month 3	(−0.7) – 6.5	(−3.8) – (−2.3)
quarter 1	19.2 – 25.1	79.2 – 80.6
end-quarter 2, cumulative	37.9 – 45.3	78.5 – 82.2
end-quarter 3, cumulative	61.2 – 74.6	78.8 – 80.4
end-year 1, cumulative	80.6 – 88.1	77.2 – 78.8
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)		
	Germany	Japan
month 1	6.5 – 10.0	15.4 – 15.7
month 2	12.6 – 14.9	43.0 – 43.8
month 3	12.3 – 13.9	(−0.7) – 0.2
quarter 1	34.2 – 36.1	58.0 – 59.3
end-quarter 2, cumulative	38.7 – 49.7	37.3 – 38.6
end-quarter 3, cumulative	62.3 – 67.1	43.2 – 46.1
end-year 1, cumulative	70.1 – 78.0	44.3 – 46.5

Table 12: VAR Estimates of the Pass-Through on Export Prices Obtained Using Import and Export Unit Values

Pass-Through on the *Consumer Price Index* Following NEER Depreciation, %			
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)			
	United States	Germany	Japan
month 1	$0.0 - 0.6$	$0.0 - 3.1$	$0.0 - 0.1$
month 2	$(-0.3) - (-0.1)$	$3.1 - 4.4$	$(-0.6) - (-0.6)$
month 3	$(-1.3) - (-1.2)$	$1.4 - 2.6$	$1.2 - 1.3$
quarter 1	$(-1.6) - (-0.7)$	$5.4 - 9.1$	$0.7 - 0.8$
end-quarter 2, cumulative	$(-3.0) - (-1.9)$	$6.5 - 11.9$	$1.9 - 2.1$
end-quarter 3, cumulative	$(-1.4) - 0.0$	$10.6 - 17.3$	$4.8 - 4.9$
year 1, cumulative	$(-1.8) - 0.3$	$15.0 - 21.4$	$6.0 - 6.2$
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)			
	United States	Germany	Japan
month 1	$0.0 - 0.7$	$0.0 - 4.0$	$(-0.4) - 0.0$
month 2	$(-0.9) - (-0.6)$	$4.8 - 5.6$	$(-1.3) - (-1.0)$
month 3	$(-2.2) - (-2.0)$	$4.2 - 4.6$	$3.1 - 3.1$
quarter 1	$(-3.1) - (-1.9)$	$9.4 - 13.8$	$1.5 - 1.9$
end-quarter 2, cumulative	$(-4.7) - (-3.2)$	$22.0 - 27.8$	$4.3 - 4.8$
end-quarter 3, cumulative	$(-1.3) - 0.9$	$29.7 - 35.9$	$7.6 - 8.2$
end-year 1, cumulative	$(-3.4) - (-0.2)$	$42.1 - 50.1$	$8.5 - 9.2$
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)			
	United States	Germany	Japan
month 1	$0.0 - 0.1$	$0.0 - 3.1$	$0.0 - 0.9$
month 2	$0.0 - 0.2$	$(-1.0) - 0.4$	$(-1.3) - (-1.1)$
month 3	$(-1.0) - (-0.8)$	$(-4.0) - (-2.0)$	$(-0.2) - (-0.1)$
quarter 1	$(-0.7) - 0.1$	$(-3.8) - 1.2$	$(-1.3) - (-0.4)$
end-quarter 2, cumulative	$(-2.0) - (-0.8)$	$(-15.4) - (-7.0)$	$(-1.0) - 0.0$
end-quarter 3, cumulative	$(-2.0) - (-0.8)$	$(-14.2) - (-6.8)$	$(-1.0) - 0.0$
end-year 1, cumulative	$0.1 - 1.3$	$(-14.8) - (-5.7)$	$(-0.5) - 0.8$

Table 13: VAR Estimates of the Pass-Through on Consumer Prices Obtained Using Import and Export Price Indices

Pass-Through on the *Consumer Price Index* Following NEER Depreciation, %		
PANEL I: *Whole* Sample Period (July 1979 - June 2002, 276 observations)		
	Germany	Japan
month 1	0.0 – 2.2	0.0 – 0.5
month 2	4.5 – 5.6	(−0.6) – (−0.6)
month 3	2.2 – 3.0	1.5 – 1.6
quarter 1	7.2 – 10.2	0.9 – 1.4
end-quarter 2, cumulative	8.6 – 12.3	2.2 – 2.7
end-quarter 3, cumulative	10.8 – 15.6	4.7 – 5.3
year 1, cumulative	15.1 – 19.7	5.9 – 6.8
PANEL II: *Early* Sample Period (July 1979 - December 1990, 138 observations)		
	Germany	Japan
month 1	0.0 – 3.5	0.0 – 1.6
month 2	3.5 – 4.9	(−2.1) – (−1.6)
month 3	6.1 – 7.5	5.0 – 5.6
quarter 1	10.4 – 14.8	3.7 – 4.8
end-quarter 2, cumulative	21.2 – 26.9	6.6 – 7.8
end-quarter 3, cumulative	27.1 – 33.2	9.2 – 11.4
end-year 1, cumulative	37.3 – 45.3	12.8 – 15.6
PANEL III: *Late* Sample Period (January 1991 - June 2002, 138 observations)		
	Germany	Japan
month 1	0.0 – 8.1	(−1.2) – 0.0
month 2	(−0.1) – 1.7	(−1.2) – (−1.2)
month 3	(−2.7) – (−1.2)	0.2 – 0.3
quarter 1	(−3.2) – 7.2	(−1.0) – (−0.5)
end-quarter 2, cumulative	(−13.0) – 1.2	(−0.8) – (−0.2)
end-quarter 3, cumulative	(−12.9) – 1.6	(−0.6) – 0.2
end-year 1, cumulative	(−13.2) – 2.8	0.0 – 1.0

Table 14: VAR Estimates of the Pass-Through on Consumer Prices Obtained Using Import and Export Unit Values

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