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# The Yen Exchange Rate and the Hollowing-out of Japanese Industry\*

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Since the demise of the Bretton Woods, the yen has seen several episodes of strong appreciation, including in the late 1970s, after the 1985 Plaza Agreement, the early and late 1990s and after 2008. These appreciations have not only been associated with “expensive yen recessions” resulting from negative effects on exports; since the late 1980s, the strong yen has also raised concerns about a de-industrialisation of the Japanese economy. Against this backdrop, the paper investigates the effects of changes to the yen exchange rate on the hollowing out of the Japanese industrial sector. To this end, the paper uses both aggregate and industry-specific data to gauge the effects of yen fluctuations on the output and exports of different Japanese industries, exploiting new data for industry-specific real effective exchange rates. Our findings support the view that the periods of yen appreciation had more than just transitory effects on Japanese manufacturing. The results also provide indication of hysteresis effects on manufacturing. While there are certainly also other factors that have contributed to a hollowing out of Japanese industry, a strong yen played a role, too.

*Keywords:* Yen appreciation, exchange rates, Japanese manufacturing, hollowing out, hysteresis.

*JEL Classification:* F31, O14.

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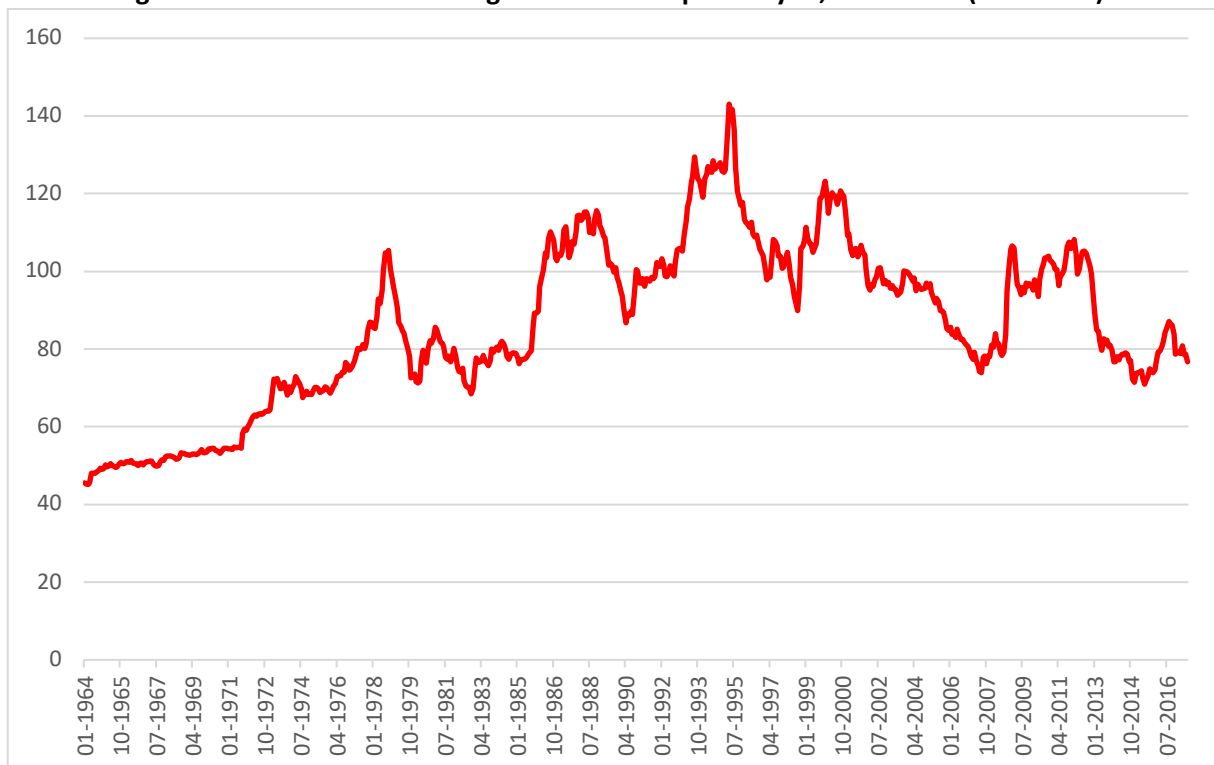
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## 1. Introduction

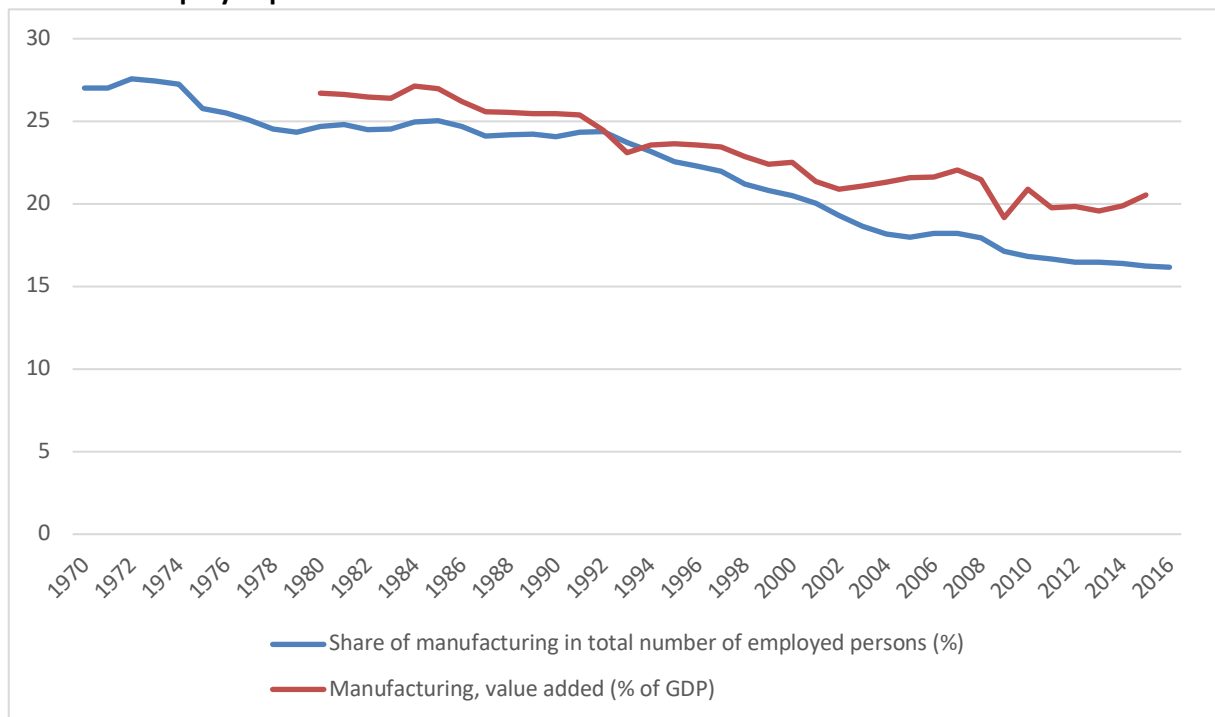
A strong yen – endaka (円高) in Japanese – has repeatedly cause distress among Japanese policymakers and manufacturers. Since the demise of the Bretton Woods system in 1971, the yen has seen several episodes of strong appreciation, including in the late 1970s, after the 1985 Plaza Agreement, the early and late 1990s and after 2008 (Figure 1). These appreciations have not only been associated with “expensive yen recessions” – endaka fukyo (円高不況) in Japanese – resulting from negative effects on exports; since the late 1980s, the strong yen has also raised concern about a de-industrialisation of the Japanese economy. Indeed, the share of manufacturing in total output has declined from 26.7% in 1980 to 20.5% in 2015, while the share of manufacturing in total number of employed persons has declined from 27.0% to 16.2% between 1970 and 2016 (Figure 2).

**Figure 1: Real effective exchange rate of the Japanese yen, 1964-2017 (2010=100)**



Source: Compiled by authors with data from the BIS Effective Exchange Rate Indices (updated 17 August 2017), [www.bis.org/statistics/eer.htm](http://www.bis.org/statistics/eer.htm).

**Figure 2: Value added in manufacturing as share of GDP and share of manufacturing in total number of employed persons**



*Source: Compiled by authors with data from the Japan Labor Force Survey and the World Development Indicators.*

While the strong yen and its potentially de-industrialising effect have received much attention in the political and economic policy discourse in Japan, there has been surprisingly little research on this in the academic literature. As pointed out by Hamada and Okada (2009: 200), most research on Japan’s “lost decade” and hollowing out of the Japanese industry “have been broadly focused on its real and domestic aspects, such as total factor productivity (TFP), growth decline, non-performing loans, and governance.” Hamada and Okada (2009: 200) argue that monetary and exchange rate policy have played an important role in this, and that “Japanese industries endured a heavy burden” due to a greatly overvalued real exchange rate.

Against this backdrop, the paper aims to investigate the effects of yen fluctuations on industrial production and the hollowing out of the Japanese industrial sector. To this end, this study conducts an econometric analysis using new data for industry-specific real effective exchange rates (RIETI, 2016) to gauge the effects of yen fluctuations on the output and exports of different Japanese industries at monthly frequency. We also use annual data since 1970 to investigate potential long-term effects of yen exchange rate changes on manufacturing employment. Following Rowthorn and Coutts (2004: 767), we define de-industrialisation – to which we also refer as “hollowing out” – as “a secular decline in the share of manufacturing in national employment”.

The topic is not only of relevance to the Japanese economy, it also bears relevance for other advanced economies. Indeed, recent years have seen an intense discussion in several member countries of the Eurozone about purported negative effects of a too high euro exchange rate on manufacturing and employment (e.g., Belke and Volz, 2015). For instance, Bénassy-Quéré et al. (2014: 7) estimate that for France “a 10% depreciation in the euro in relation to a partner country outside the eurozone increases the value of the average exporter’s sales to this country by around 5-6%”, with most of this effect

realised in the same year as the depreciation. Likewise, the US has recently seen a renewed discussion on de-industrialisation, with a prominent role of the dollar exchange rate in this discourse (e.g., Campbell, 2017; Levinson, 2017). A better understanding of any exchange rate effects on Japanese manufacturing should also provide insights for economic and monetary policy elsewhere.

The remainder of the paper is structured as follows. Section 2 provides a review of the literature on the effects of exchange rate changes on de-industrialisation. Section 3 comprises our empirical analysis. The empirical part is split into two sections: Firstly, we conduct an empirical analysis with annual, aggregate data from 1970 to 2016 that allows us to investigate potential long-term effects of yen exchange rate changes on manufacturing employment. Secondly, we use monthly, industry-specific data from 2001:M01 to 2017:M06 to gauge the dynamics of yen exchange rate changes on Japanese sectoral manufacturing employment. Section 4 concludes.

## 2. Literature review

It is well established that depreciated real exchange rates can help to stimulate industrial development and economic growth (Rodrik, 2008). Although there seems to exist a “fear of appreciation” (Levy-Yeyati et al., 2013) relatively little research has been conducted on the effect of overvalued exchange rates on economies. The first systematic empirical cross-country analysis to investigate the effects of large exchange rate appreciation on current account balances and on real output was conducted by Kappler et al. (2013), who find that episodes of strong exchange rate appreciations are associated with deteriorating current account balances and a slow-down of real export growth, but no significant effects on output. Likewise, Bussière et al. (2015) examine to what extent large and rapid real exchange rate appreciations impact on economic growth. Using a sample of 53 emerging and advanced economies they find that while large appreciations dampen export growth and boost import growth, output growth is higher on average.

The literature on de-industrialisation has focused less on exchange rate valuations than on changes in specialisation, consumption, technological progress and productivity, international trade and investment patterns (Rowthorn and Coutts, 2004). The major exemption is the US economy, for which the effect of the dollar exchange rate on the US industry has been studied widely. The findings in this literature are mixed, with some studies finding a large effect (Branson and Love, 1986, 1987, 1988a, 1988b; Revenga, 1992; Gourinchas, 1999; Kandil and Mirzaie, 2002, 2003; Klein et al., 2003; Campbell, 2017) and others a small or no impact (Glick and Hutchison, 1990; Goldberg, 1993; Campa and Goldberg, 1995 and 2001; Bahmani-Oskooee et al., 2000, 2007).<sup>1</sup>

According to Rowthorn and Ramaswamy (1997a, 1997b) and Rowthorn and Coutts (2004, 2013), the most important internal/domestic factors contributing to deindustrialization are productivity growth in the manufacturing sector and shifting patterns of domestic expenditure, i.e., with rising income consumptions tends to shifts towards more services and away from manufactured goods (Bell’s Law).

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<sup>1</sup> There are also a number of studies investigating exchange rate effects on European economies. For instance, using firm-level data, Moser et al. (2010) and Berman et al. (2012) find relatively small exchange rate effects on employment of German and French firms, respectively. Ekholm et al. (2012) analyse the impact of a real appreciation of the Norwegian krona on Norwegian net-exporting and import-competing firms in the early 2000s, finding negative effects on domestic employment across both groups of firms. Thorbecke and Kato (2017) examine the effect of changes to the Swiss franc exchange rate on the Swiss economy and find differences related to the sophistication of exports.

The most important external factors are identified as “North-South” trade and the outsourcing of labour-intensive manufacturing to low-wage countries. However, domestic factors, especially a higher productivity growth in manufacturing than in services, are widely seen as the dominant factor behind deindustrialization (Rowthorn and Ramaswamy, 1997; Rowthorn and Coutts, 2004, 2013).<sup>2</sup>

Surprisingly few studies have looked into the potentially de-industrialising effect of endeka, even though Obstfeld (2010) pointed out that Japan’s real economic growth rate has been strongly negatively correlated with the level of the yen’s real effective exchange rate. Only few studies can be found that systematically try to verify the hypothesis that the strong yen has contributed to de-industrialisation and outsourcing of industry in Japan, including Dekle (1996), Dekle et al. (2010) and Yamashita (2013, 2015). Dekle (1996) and Dekle et al. (2010) find that yen appreciations of 1985 and 1995 significantly hurt the ability of Japan to compete with the U.S. by raising the relative production costs of Japanese industries. Yamashita (2013, 2015) finds that yen appreciation and outward FDI contributed to deindustrialization and suggests that a weak yen may contribute to reshoring.

A string of studies examines the role of Japanese outward direct investment in the hollowing out process. Japanese outward direct investment took off in the mid-1980s (Fukao et al., 2003). The number of Japanese multinationals increased by 290% between 1985 and 1992, while overseas production of Japanese firms increased from 3% in 1982 to 17% in 2002 (Ryan and Toubal, 2017). Analysing data on Japanese acquisitions in the United States from 1975-1992, Blonigen (1997) finds that real dollar depreciations increased the likelihood of Japanese acquisitions in U.S. industries. Likewise, the BOJ (2011) attributed the renewed shifting of production abroad since 2008 to the relative appreciation of the yen compared to the levels before the Lehman shock.

Empirically, the effect of outsourcing on domestic manufacturing employment is mixed. According to Baldwin (2006), offshoring labour-intensive production processes to adjacent lower-cost locations in Southeast and Northeast Asia since the mid-1980s did apparently allow Japanese companies to retain parts of manufacturing production domestically while maintaining international competitiveness. The BOJ (1989) finds that increased outward direct investment contributed to long-term structural adjustment in the export sector. This contributed to a temporary surge in the export of capital goods and parts to overseas production bases of Japanese companies. Yamashita and Fukao (2010) find no indication that expansion of overseas operations by Japanese manufacturing multinational enterprises reduced home employment over the period 1991-2002; indeed, their analysis suggests that it may have contributed to maintaining the level of home employment. A recent study by Ryan and Toubal (2017) uses Japanese firm-level data for the period 1982-2001 to compare employment between firms with and without foreign affiliates and the extent to which the development of foreign operations affects firm-level domestic employment. It finds “some limited evidence” of the hollowing out of the Japanese economy by Japanese multinational firms shifting production abroad from 1991 onwards.

Using the same industry-specific exchange rates that will be employed in our empirical analysis later, Sato et al. (2013) analyse export price competitiveness and export performance of Japan, China and Korea and find that the strong yen after 2007 destroyed the Japanese electronics sector. Likewise, Thorbecke (2012) finds in a study focused on the Japanese electronics industry that the yen

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<sup>2</sup> A further factor that appears to have played a role in the falling share of manufacturing is the outsourcing of various services that historically were conducted in-house and hence counted as manufacturing, whereas they are now counted as services (cf. Boulhol and Fontagné).

appreciation between 2007 and 2011 had a very damaging effect on electronics exports. However, neither study controls for employment effects.

An important issue in this context is that exchange rate appreciations and depreciations may not have symmetrical effects. That is, a period of strong appreciation may lead to outsourcing and the loss of manufacturing capacities – and employment – in the home country, but a depreciation later on may not bring about a return (or “re-shoring”) of manufacturing activities and employment, given that long-term investments have been made abroad, which will not be reversed. These hysteresis effects have been described in sunk cost models of trade hysteresis (Krugman, 1988; Baldwin and Krugman, 1989; Baldwin, 1990; Baldwin and Lyons, 1994) and empirically investigated for the U.S. and a number of European economies.<sup>3</sup>

### 3. Empirical approach and findings

For our econometric analysis, we use both aggregate and industry-specific data to gauge the effects of changes of the real effective yen exchange rate. Taking account of the availability of different data series, we first conduct an empirical analysis with annual data from 1970 to 2016 that allows us to investigate potential long-term effects of yen exchange rate changes on manufacturing employment. Secondly, we use monthly data from 2001:M01 to 2017:M06 to check the dynamics of yen exchange rate changes on Japanese sectoral manufacturing employment, exploiting a new quarterly dataset on Japanese sectoral exchange rates which are only available for this time span. For this purpose, we use a mixed time series-panel framework.

#### 3.1 Analysis with annual data, 1970-2016

##### 3.1.1 Data and empirical model

Building on Rowthorn and Ramaswamy (1997a) and Rowthorn and Coutts (2004, 2013), we estimate the following model for Japan:<sup>4</sup>

$$(1) \text{ EMPMAN} = \beta_0 + \beta_1 \text{ EMPMAN}_{t-1} + \beta_2 \text{ CAP} + \beta_3 \text{ EXPMAN} + \beta_4 \text{ OFDI} + \beta_5 \text{ TFP} + \beta_6 \text{ EXPCHIN} + \beta_7 \text{ EXR} + \varepsilon,$$

with *EMPMAN* = Share of manufacturing in total number of employed persons (%), *CAP* = Gross domestic fixed capital formation (% of GDP), *EXPMAN* = Manufacturing exports, *OFDI* = Outward direct investment, *TFP* = Total factor productivity at constant national prices, *EXPCHIN* = Chinese exports, and *EXR* = Real effective exchange rate. Appendix 1 provides details on the variables and sources.

Increases in fixed capital formation could theoretically have both positive and negative effects on manufacturing employment. On the one hand, an increased investment in the domestic manufacturing sector could have positive employment effects. On the other hand, however, capital investment could advance the automation of manufacturing production, with neutral or even negative effects on manufacturing employment. In contrast, an increase in manufacturing exports should clearly have a

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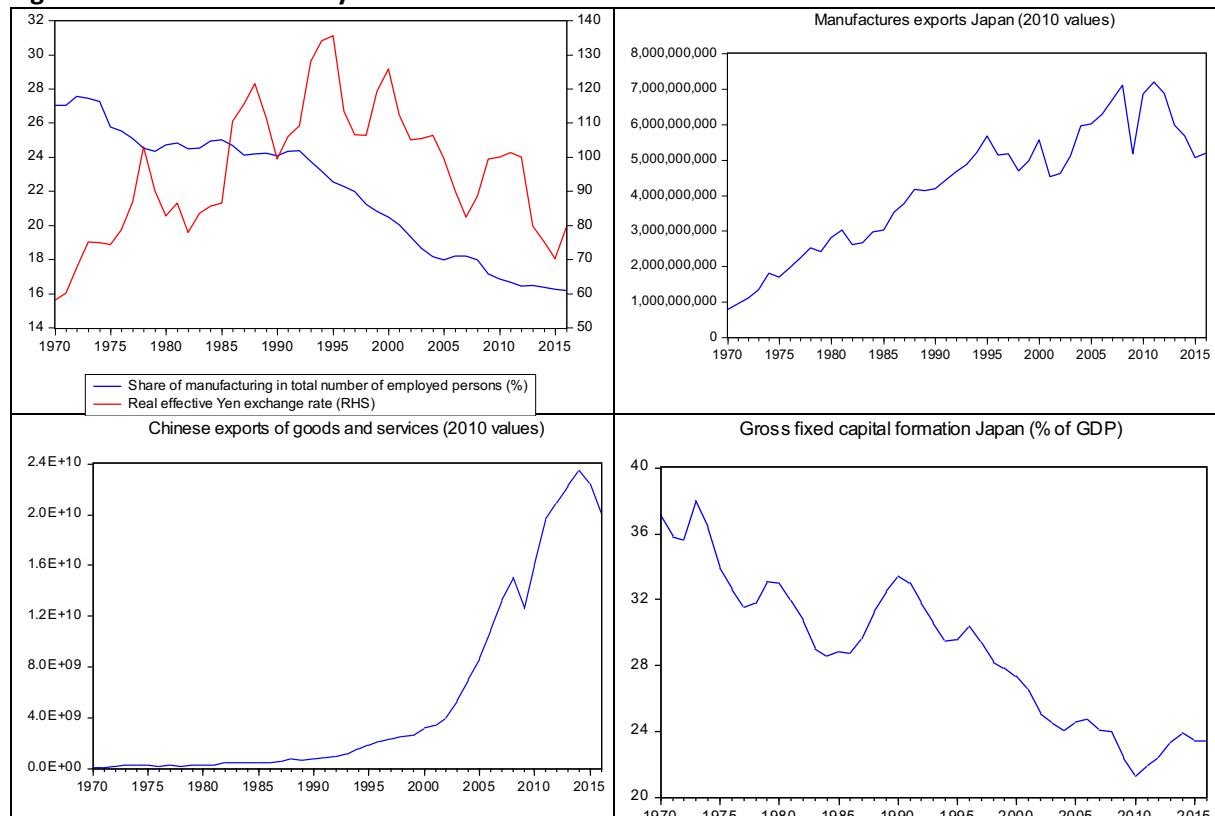
<sup>3</sup> Cf. Belke et al. (2013), Belke and Kronen (2016), Campa (2004), Campbell (2017).

<sup>4</sup> The same framework is applied by Boulhol and Fontagné (2006), who analyse the change in the share of manufacturing in employment between 1970-2002 for 16 OECD countries. None of these studies controls for exchange rate effects.

positive effect on domestic manufacturing employment. As discussed above, the effect of outward direct investment can be ambiguous: on the one hand a shifting of production abroad may have negative employment effects, but at the same time there could be positive outcomes because it may increase the competitiveness of domestic operations which would remain important parts of regional or global value chains (Baldwin, 2006; Yamashita and Fukao, 2010). As mentioned above, TFP improvements have been identified in the de-industrialisation literature as a major factor, so we would expect a negative coefficient. The effects Chinese export growth are, again, ambiguous.<sup>5</sup> While Chinese export growth has been a reflection of China's rapid growth process, which has also opened up great opportunities for Japanese companies, Chinese exporters have increasingly become competitors of Japanese manufacturing firms. Last but not least, we would expect the coefficient estimate for the real effective exchange rate – our main variable of interest – to be negative if the hollowing out hypothesis holds.

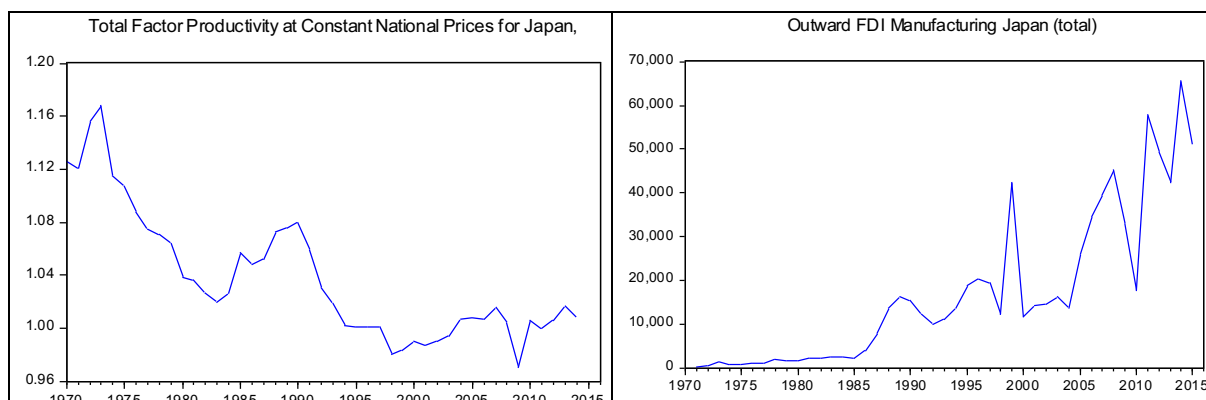
Figure 3 graphically displays the time series of the variables used in our estimations based on annual data. It conveys a first impression of the stationarity properties of the time series used. At first glance, all time series appear to contain, probably stochastic, trends. We will explicitly check for stochastic trends before we start our estimation exercise in this section in order to make sure that the stochastic properties of the included variables meet the standard assumptions of our regression analysis. Moreover, some first indications of correlations can be derived as well. Whether this means causation in an econometric sense will be checked later on in this section.

**Figure 3: Variables for analysis with annual data**



<sup>5</sup> Unfortunately, no data was available for Chinese manufacturing exports for this period, which is why we selected total exports from China.





Source: Compiled by authors.

The top left panel of Figure 3 shows the two main variables of interest, the share of manufacturing in total number of employed persons and the real effective exchange rate. Between 1970 and the mid-1990s, the real effective exchange rate showed a clear long-term appreciation trend, but with two periods – 1978-1982 and 1988-1990 – of depreciation. After a peak in 1995, the real effective exchange rate then shows a declining trend, albeit with periods of strong appreciation in 1998-2000, 2007-2011 and 2015-2016. Over the entire sample period, the share of manufacturing in total number of employed persons shows a long-term declining trend, although the share of manufacturing remains stable over the periods 1987-1992, 2005-2007 and 2012-2016.

We continue with a brief look at the ordinary Pearson correlations among the variables included in our regressions (Table 1). The purpose of it is twofold. First, a brief inspection of the correlation pattern may serve as a first hint at potential multicollinearities present in our regressions. And second, it may reveal first evidence on the expected sign of the impact of one variable of interest (e.g., the yen real effective exchange rate, EXR) on another (e.g., Japanese employment in manufacturing EMPMAN). For instance, the yen real effective exchange rate appears to be negatively correlated with the share of manufacturing in total number of employed persons EMPMAN (a relation which is at the centre of interest in this paper) and gross domestic fixed capital formation CAP. It is positively correlated with outward direct investment OFDI. At first glance, these correlations are all in line with our model.

**Table 1: Ordinary Pearson correlations**

	CAP	EMPMAN	EXPCHIN	EXPMAN	EXR	OFDI	TFP
CAP	1.000000	0.946907	-0.796667	-0.846322	-0.239863	-0.730577	0.807738
EMPMAN	0.946907	1.000000	-0.868595	-0.890158	-0.216322	-0.833582	0.762477
EXPCHIN	-0.796667	-0.868595	1.000000	0.726566	-0.114324	0.879415	-0.450889
EXPMAN	-0.846322	-0.890158	0.726566	1.000000	0.496410	0.796646	-0.783812
EXR	-0.239863	-0.216322	-0.114324	0.496410	1.000000	0.116064	-0.531790
OFDI	-0.730577	-0.833582	0.879415	0.796646	0.116064	1.000000	-0.545285
TFP	0.807738	0.762477	-0.450889	-0.783812	-0.531790	-0.545285	1.000000

Source: Compiled by authors.

However, the correlation table also indicates a positive correlation of the exchange rate with manufacturing exports EXPMAN. Regarding our regression analysis two issues emerging from this specific finding come to mind. Firstly, there is a certain agreement in the literature that one may well find impacts of exchange rate changes on employment even if there is less evidence of an impact on trade (Belke and Gros, 2001). Secondly, a thorough regression analysis is necessary to take into account

leads and lags of the variables when it comes to establish any one-way “causation”. This is dramatically shown by the main results of the export hysteresis literature which shows that a positive coefficient of the exchange rate in export, employment, investment or industrial output equations may just mirror and hide a more complex dynamic of the relationship which may include stepwise negative effects of an appreciation of the home currency (Krugman and Baldwin, 1987; Belke et al., 2013).<sup>6</sup>

However, we would like to stress again at this early stage of analysis that correlation does not necessarily mean establishing causation in an econometric sense. We leave this important task to the remainder of this section.

### 3.1.2 Unit root tests

We now proceed with unit root tests of the variables to be employed in our empirical model. The results of our single time series unit root tests are displayed in Table 2. According to Table 2, all variables turn out to be integrated of order one (I(1)), except Chinese exports EXPCHIN which appears borderline I(2).

**Table 2: Results of unit root tests (ADF test statistics and probabilities)**

Variable	Level	First difference	Integration order of variable
CAP	-1.33 (0.6087)	-4.96 (0.0002)	I(1)
EMPMAN	-0.56 (0.8680)	-4.69 (0.0004)	I(1)
EXPCHIN	2.52 (1.0000)	-0.59 (0.9735)	I(1)/I(2)
EXPMAN	-1.92 (0.3195)	-8.10 (0.0000)	I(1)
EXR	-2.75 (0.0088)	-5.20 (0.0001)	I(1)
OFDI	0.45 (0.9827)	-8.23 (0.0000)	I(1)
TFP	-1.83 (0.3629)	-6.17 (0.0000)	I(1)

*Source: Compiled by authors.*

*Note: Table displays empirical realisations of the Augmented Dickey-Fuller test with a constant (MacKinnon (1996) one-sided p-values).*

For the real effective yen exchange rate EXR, all complementary unit root tests we conducted beyond the tabulated ADF-test (such as the Dickey-Fuller GLS (ERS) test, the Phillips-Perron test, the Kwiatkowski-Phillips-Schmidt-Shin test and the Ng-Perron unit root test) except the AFD-test point into the direction of an I(1) process. Hence, we proceed with the yen exchange rate considering it as an I(1) variable. The results are available on request.

As a robustness check, we also conducted Dickey-Fuller GLS tests, Phillips-Perron tests, Kwiatkowski-Phillips-Schmidt-Shin (KPSS) and Ng-Perron unit root tests. The results overall stayed the same. Among others, they support treating EXPCHIN finally as an I(1) variable in our regressions.

Given the unit root tests conducted in this section and their relatively unambiguous results, we feel legitimised to employ the levels of the variables in the framework of a cointegration approach (option 1) and/or use the first differences of all variables as stationary variables in our estimations (option 2).

<sup>6</sup> The hysteresis argument is mainly related to the irreversibility of investment-type decisions, due to high fixed costs which cannot be recovered ex post (sunk costs). Industrial production, the focus of this paper, is a very good example because it is dependent on investment-type decisions related to, among others, fixed investment, hiring and firing decisions and establishing distribution networks.

We now turn to the first option, the estimation of a cointegrating equation in order to assess the impact of the yen exchange rate (EXR) on the share of manufacturing in total number of employed persons in Japan (EMPMAN), taking the other variables contained in eq. (1) into account simultaneously in the long-run relationship.

### 3.1.3 A cointegration approach: in search of long-run “causation” of the “hollowing out”

As our cointegration framework we choose Dynamic Ordinary Least Squares (DOLS). This is because DOLS allows us to test for *unequivocal causality* running from the independent to the dependent variables. The DOLS estimation technique *controls for endogeneity* of explanatory variables (Stock and Watson, 1993; Wooldridge, 2009). Endogeneity in the form of feedback effects or reverse causality between the dependent and independent variables would lead to a misspecification of our estimated models. More specifically, the DOLS procedure controls for endogeneity of all explanatory variables by inserting *leads and lags* of the changes of all exogenous variables. Hence, DOLS is a *powerful estimation technique*, where standard errors are corrected for heteroscedasticity and cross-section correlation (Saikkonen, 1991; Stock and Watson, 1993). By inserting the leads and lags of the exogenous variables in first differences, the explanatory variables in levels become (*super-*) *exogenous* and the regression results *unbiased* (Wooldridge, 2009).

Table 3 displays the results of a DOLS estimation of our fully specified empirical model (eq. (1)). The share of manufacturing in total number of employed persons (EMPMAN) represents the dependent variable. The estimated coefficient of the yen exchange rate (marked in bold letters in Tab. 3) is negative as expected from theory and highly significant. Seen on the whole, also the other estimated coefficients are significant and their signs corresponding with our priors. TFP shows a negative coefficient, suggesting that higher productivity has resulted in a lower share of manufacturing employment, but this effect is not statistically significant. Gross domestic fixed capital formation shows a positive sign and is highly significant. As expected, manufacturing exports have a positive and significant effect on manufacturing, whereas outward direct investment has a negative and significant effect. Chinese exports show a positive sign, suggesting that the positive effects of Chinese economic development on Japanese manufacturing dominated over the negative competition effect. Finally, the empirical realisations of the goodness-of-fit criteria, among them the very high R-Squared, indicate the appropriateness of our selected empirical model.

**Table 3: Dynamic least squares estimation for fully specified model**

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Dependent Variable: EMPMAN  
Method: Dynamic Least Squares (DOLS)  
Sample (adjusted): 1973-2013  
Included observations: 41 after adjustments  
Cointegrating equation deterministics: C  
Fixed leads and lags specification (lead=1, lag=1)  
Long-run variance estimate (Bartlett kernel, Newey-West fixed bandwidth = 4.0000)

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Variable	Coefficient	Std. Error	t-Statistic	Prob.
TFP	-10.49578	6.642659	-1.580057	0.1337
CAP	1.108099	0.135792	8.160269	0.0000
EXPCHIN	2.20E-10	8.54E-11	2.572967	0.0204
EMPMAN	1.42E-09	3.67E-10	3.879566	0.0013

OFDI	-0.000192	3.60E-05	-5.325351	0.0001
<b>EXR</b>	<b>-0.051424</b>	<b>0.016242</b>	<b>-3.166034</b>	<b>0.0060</b>
C	2.783452	5.976907	0.465701	0.6477
<hr/>				
R-squared	0.992483	Mean dependent var	22.11307	
Adjusted R-squared	0.981208	S.D. dependent var	3.295862	
S.E. of regression	0.451804	Sum squared resid	3.266036	
Long-run variance	0.175494			

*Source: Compiled by authors.*

Already the simple Pearson correlations in Tale 1 suggest that EXPCHIN and EXPMAN may be highly multicollinear. Moreover, our robustness checks reveal that the total goodness-of-fit does not become significantly lower if EXPCHIN is eliminated from our empirical model. This does not come as a surprise since both variables represent indicators of the world business cycle, i.e., the so-called “global factor”. For this reason, we leave out EXPCHIN in our next DOLS-specification. The corresponding regression results are displayed in Table 4. Again, the estimation results clearly are in accordance with our priors.

**Table 4: Dynamic least squares estimation for model excluding Chinese exports**

Dependent Variable: EMPMAN				
Method: Dynamic Least Squares (DOLS)				
Sample (adjusted): 1973-2013				
Included observations: 41 after adjustments				
Cointegrating equation deterministics: C				
Fixed leads and lags specification (lead=1, lag=1)				
Long-run variance estimate (Bartlett kernel, Newey-West fixed bandwidth = 4.0000)				
<b>Variable</b>	<b>Coefficient</b>	<b>Std. Error</b>	<b>t-Statistic</b>	<b>Prob.</b>
TFP	-4.949611	8.354137	-0.592474	0.5602
CAP	0.930860	0.140335	6.633109	0.0000
EXPMAN	1.10E-09	4.76E-10	2.309287	0.0317
OFDI	-0.000122	3.54E-05	-3.440251	0.0026
<b>EXR</b>	<b>-0.055554</b>	<b>0.020831</b>	<b>-2.666873</b>	<b>0.0148</b>
C	3.580114	8.186393	0.437325	0.6666
<hr/>				
R-squared	0.988524	Mean dependent var	22.11307	
Adjusted R-squared	0.977048	S.D. dependent var	3.295862	
S.E. of regression	0.499324	Sum squared resid	4.986488	
Long-run variance	0.337452			

*Source: Compiled by authors.*

Tables 5 and 6 display our DOLS regression results with White correction of the coefficient covariance matrix. That is, we estimate the coefficient covariance using the White covariance (White, 1980). With this, we make use of a heteroskedasticity consistent covariance matrix estimator which provides consistent estimates of the coefficient covariances in the presence of (conditional) heteroskedasticity of unknown form. There is no substantial change of the estimation results, with respect to the magnitude and the sign of the estimated coefficients. If at all, the significance of, for instance, the yen exchange rate, even slightly increases once again.

**Table 5: Dynamic least squares estimation of fully specified model with White correction**

Dependent Variable: EMPMAN  
Method: Dynamic Least Squares (DOLS)  
Sample (adjusted): 1973-2013  
Included observations: 41 after adjustments  
Cointegrating equation deterministics: C  
Fixed leads and lags specification (lead=1, lag=1)  
White heteroskedasticity-consistent standard errors & covariance

Variable	Coefficient	Std. Error	t-Statistic	Prob.
TFP	-10.49578	6.116778	-1.715900	0.1055
CAP	1.108099	0.114299	9.694747	0.0000
EXPCHIN	2.20E-10	8.05E-11	2.728241	0.0149
EXPMAN	1.42E-09	3.44E-10	4.133799	0.0008
OFDI	-0.000192	3.68E-05	-5.205931	0.0001
<b>EXR</b>	<b>-0.051424</b>	<b>0.016735</b>	<b>-3.072769</b>	<b>0.0073</b>
C	2.783452	6.037110	0.461057	0.6510
R-squared	0.992483	Mean dependent var		22.11307
Adjusted R-squared	0.981208	S.D. dependent var		3.295862
S.E. of regression	0.451804	Sum squared resid		3.266036

Source: Compiled by authors.

**Table 6: Dynamic least squares estimation of model excluding Chinese exports with White correction**

Dependent Variable: EMPMAN  
Method: Dynamic Least Squares (DOLS)  
Sample (adjusted): 1973-2013  
Included observations: 41 after adjustments  
Cointegrating equation deterministics: C  
Fixed leads and lags specification (lead=1, lag=1)  
White heteroskedasticity-consistent standard errors & covariance

Variable	Coefficient	Std. Error	t-Statistic	Prob.
TFP	-4.949611	6.434338	-0.769249	0.4507
CAP	0.930860	0.101876	9.137229	0.0000
EXPMAN	1.10E-09	3.70E-10	2.974161	0.0075
OFDI	-0.000122	3.25E-05	-3.746083	0.0013
<b>EXR</b>	<b>-0.055554</b>	<b>0.014164</b>	<b>-3.922160</b>	<b>0.0008</b>
C	3.580114	6.497678	0.550984	0.5877
R-squared	0.988524	Mean dependent var		22.11307
Adjusted R-squared	0.977048	S.D. dependent var		3.295862
S.E. of regression	0.499324	Sum squared resid		4.986488

Source: Compiled by authors.

Tables 7 and 8 display our DOLS regression results with Newey-West correction of the coefficient covariance matrix (Newey and West, 1987). In this exercise, we apply a covariance estimator that is

consistent in the presence of both heteroskedasticity and autocorrelation (HAC) of unknown form, under the assumption that the autocorrelations between distant observations die out. In other words, we employ a sandwich-style HAC (Newey-West) covariance matrix estimator for reasons of comparison. We thus go beyond just using the White covariance matrix (as in Tables 5 and 6) which assumes that the residuals of the estimated equation are serially uncorrelated. Again, the estimation results do not change much with respect to both the magnitude and the sign of the estimated coefficients.

**Table 7: Dynamic least squares estimation of fully specified model with Newey-West correction**

Dependent Variable: EMPMAN  
Method: Dynamic Least Squares (DOLS)  
Sample (adjusted): 1973-2013  
Included observations: 41 after adjustments  
Cointegrating equation deterministics: C  
Fixed leads and lags specification (lead=1, lag=1)  
HAC standard errors & covariance (Bartlett kernel, Newey-West fixed bandwidth = 4.0000)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
TFP	-10.49578	5.232113	-2.006031	0.0621
CAP	1.108099	0.110067	10.06753	0.0000
EXPCHIN	2.20E-10	9.38E-11	2.343780	0.0323
EXPMAN	1.42E-09	3.24E-10	4.395452	0.0005
OFDI	-0.000192	3.71E-05	-5.161301	0.0001
<b>EXR</b>	<b>-0.051424</b>	<b>0.015851</b>	<b>-3.244239</b>	<b>0.0051</b>
C	2.783452	5.467176	0.509121	0.6176
R-squared	0.992483	Mean dependent var		22.11307
Adjusted R-squared	0.981208	S.D. dependent var		3.295862
S.E. of regression	0.451804	Sum squared resid		3.266036

Source: Compiled by authors.

**Table 8: Dynamic least squares estimation model excluding Chinese exports with Newey-West correction**

Dependent Variable: EMPMAN  
Method: Dynamic Least Squares (DOLS)  
Sample (adjusted): 1973-2013  
Included observations: 41 after adjustments  
Cointegrating equation deterministics: C  
Fixed leads and lags specification (lead=1, lag=1)  
HAC standard errors & covariance (Bartlett kernel, Newey-West fixed bandwidth = 4.0000)

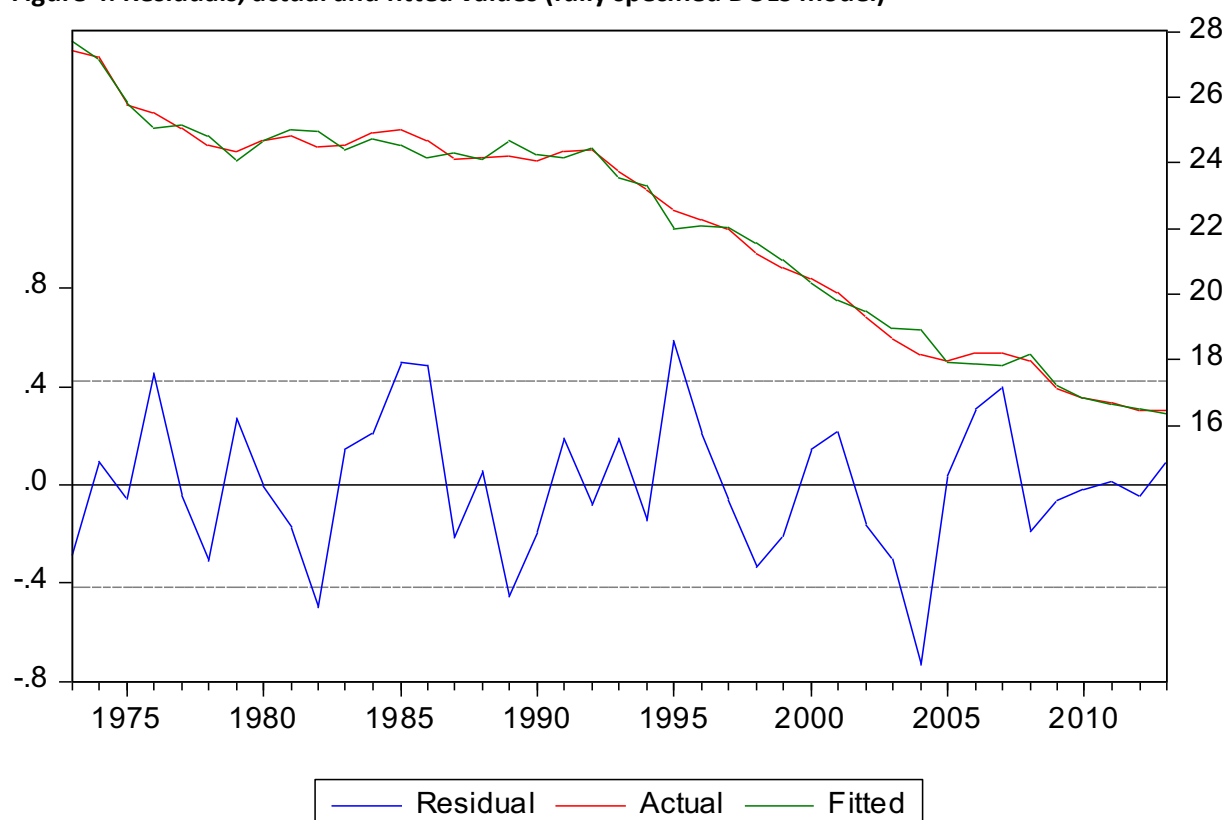
Variable	Coefficient	Std. Error	t-Statistic	Prob.
TFP	-4.949611	7.467324	-0.662836	0.5150
CAP	0.930860	0.100093	9.299968	0.0000
EXPMAN	1.10E-09	4.34E-10	2.532957	0.0198
OFDI	-0.000122	4.22E-05	-2.881250	0.0092
<b>EXR</b>	<b>-0.055554</b>	<b>0.011886</b>	<b>-4.673997</b>	<b>0.0001</b>
C	3.580114	7.497427	0.477512	0.6382
R-squared	0.988524	Mean dependent var		22.11307

Adjusted R-squared	0.977048	S.D. dependent var	3.295862
S.E. of regression	0.499324	Sum squared resid	4.986488

Source: Compiled by authors.

Application of the DOLS procedure requires time series to be non-stationary *and* to be cointegrated over time. We are able to show that both conditions are satisfied in our case. Evidence of non-stationarity has been delivered by the results of our unit root tests in section 3.1.2 and empirical support of cointegration among the variables contained in our empirical models is conveyed by cointegration tests (Table 9 for the fully specified model and Table 10 for the model without Chinese exports). Figures 4 and 5 display the residuals (which form the basis for the Engle-Granger cointegration tests described in Tables 9 and 10) and the actual and fitted values according to the regression models described in Tables 3 and 4, respectively.

**Figure 4: Residuals, actual and fitted values (fully specified DOLS model)**



Source: Compiled by authors.

**Table 9: Test for cointegration (for fully specified DOLS model)**

Cointegration Test - Engle-Granger  
Equation: EQ01  
Specification: EMPMAN TFP CAP EXPCHIN EXPMAN OFDI EXR C  
Cointegrating equation deterministics: C  
Null hypothesis: Series are not cointegrated

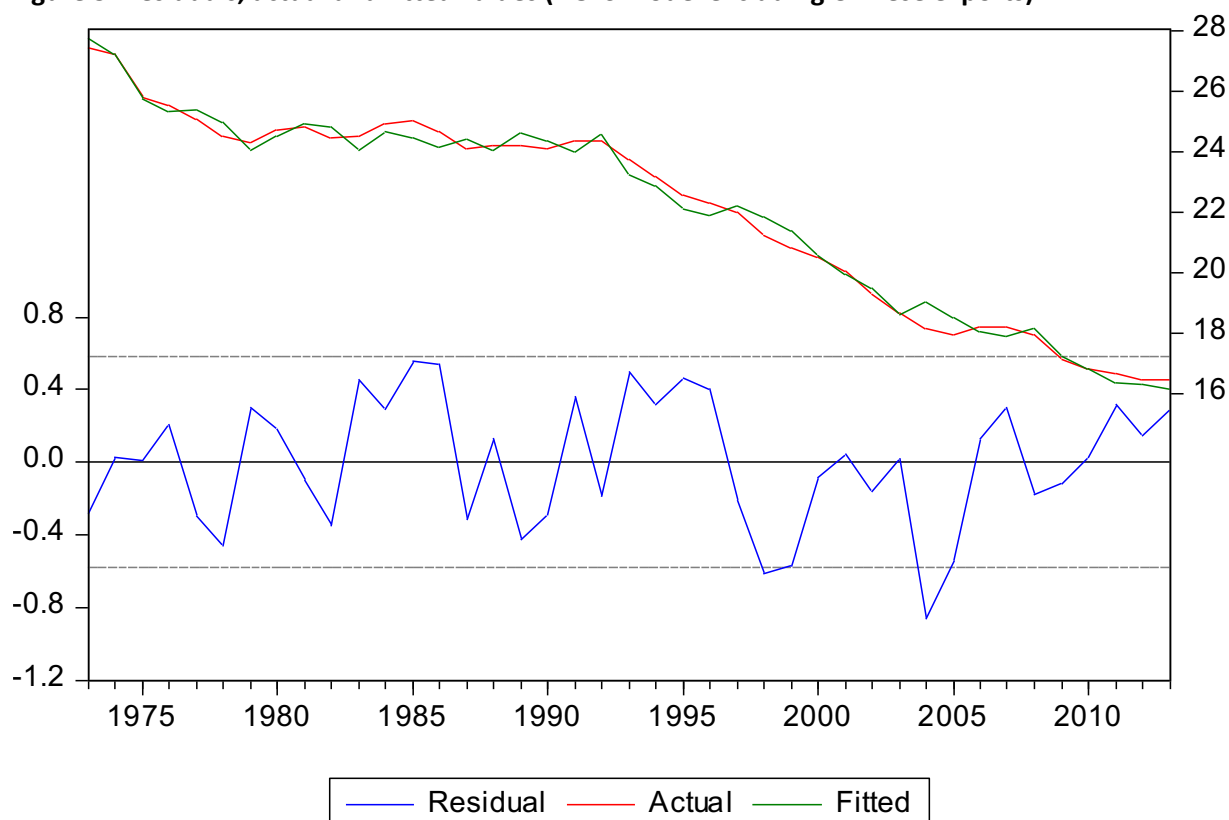
Engle-Granger Test Equation:

Dependent Variable: D(RESID)  
 Method: Least Squares  
 Date: 10/20/17 Time: 16:34  
 Sample (adjusted): 1972-2014  
 Included observations: 43 after adjustments

Variable	Coefficient	Std. Error	t-Statistic	Prob.
RESID(-1)	-0.282619	0.107162	-2.637302	0.0117
R-squared	0.141977	Mean dependent var		0.005944
Adjusted R-squared	0.141977	S.D. dependent var		0.559905
S.E. of regression	0.518638	Akaike info criterion		1.547760
Sum squared resid	11.29738	Schwarz criterion		1.588718
Log likelihood	-32.27683	Hannan-Quinn criter.		1.562864
Durbin-Watson stat	1.855124			

Source: Compiled by authors.

Figure 5: Residuals, actual and fitted values (DOLS model excluding Chinese exports)



Source: Compiled by authors.

Table 10: Test for cointegration (for DOLS model without Chinese exports)

Cointegration Test - Engle-Granger  
 Equation: EQ01  
 Specification: EMPMAN TFP CAP EXPMAN OFDI EXR C  
 Cointegrating equation deterministic: C  
 Null hypothesis: Series are not cointegrated  
 Engle-Granger Test Equation:  
 Dependent Variable: D(RESID)  
 Method: Least Squares



Sample (adjusted): 1972-2014  
 Included observations: 43 after adjustments

Variable	Coefficient	Std. Error	t-Statistic	Prob.
RESID(-1)	-0.367846	0.119751	-3.071757	0.0037
R-squared	0.183441	Mean dependent var		-0.001782
Adjusted R-squared	0.183441	S.D. dependent var		0.677695
S.E. of regression	0.612390	Akaike info criterion		1.880087
Sum squared resid	15.75091	Schwarz criterion		1.921045
Log likelihood	-39.42186	Hannan-Quinn criter.		1.895191
Durbin-Watson stat	1.861195			

Source: Compiled by authors.

It turns out, that the residuals in Figures 4 and 5 are well-behaved, i.e., stationary, and that the Engle-Granger tests do reject the null hypothesis of no cointegration. Seen on the whole, thus, we feel legitimised to argue that the yen real effective exchange rate has enacted a negative impact on the share of manufacturing in total number of employed persons in Japan over the long run from 1970 to 2016.

We now turn to the second option mentioned above, that is to conduct Autoregressive Distributed Lag (ARDL) estimations based on the first differences of the non-stationary variables.

### 3.1.4 ARDL estimations: in search of the dynamics underlying the “hollowing out”

In this section, we employ Autoregressive Distributed Lag (ARDL) estimations in order to estimate the intertemporal dynamics. ARDLs are standard least squares regressions that include lags of both the dependent variable and explanatory variables as regressors (Greene, 2008). Their application requires the data to be stationary. This leads us to employ (only) the first differences of the variables of our empirical model. This comes at the cost that we are not able to exploit level information (as it is the case in our cointegration exercise). The ARDL model selection process employed by us uses the same sample for each estimation and selects the final model according to the empirical realisation of information criteria (in our case, Akaike). The empirical results of the ARDL exercise is displayed in Table 11 below.

The expression ARDL(1, 2, 0, 1, 1, 0, 0) in Table 11 indicates the model by our search algorithm selected according to the Akaike information criterion. The first entry in brackets denotes the lags of the endogenous lagged variable. The second entry corresponds with the lag number of the change of total factor productivity, the third entry indicates the lag number of the change in gross domestic fixed capital formation (CAP). The fourth one stands for the lag number of the change in manufacturing exports (EXPMAN). The fifth entry denotes the lag number of the change in Chinese exports (EXPCHIN). The sixth entry is related to the lag number of the outward direct investment variable (OFDI). And the seventh entry refers to the lag number of lags of the yen exchange rate (EXR). The result column containing the yen exchange rate effect denotes the significant estimated coefficient for the exchange rate (-0.016393) and its significance level (0.02) in the selected ARDL model. Values in brackets in Table 11 denote the respective significance level (probabilities). Once again, we find a positive and significant effect of the yen exchange rate on the share of manufacturing employment.

**Table 11: ARDL estimation of fully specified model**

Variable	Coefficient	Std. Error	t-Statistic	Prob.*
D(EMPMAN(-1))	0.391195	0.141019	2.774061	0.0094
D(TFP)	-3.808531	4.248908	-0.896355	0.3772
D(TFP(-1))	2.768530	3.802030	0.728171	0.4722
D(TFP(-2))	-6.491839	3.288327	-1.974207	0.0576
D(CAP)	0.109411	0.071213	1.536395	0.1349
D(EXPMAN)	4.35E-10	1.70E-10	2.564722	0.0156
D(EXPMAN(-1))	-2.45E-10	1.18E-10	-2.081662	0.0460
D(EXPCHIN)	-1.47E-10	1.01E-10	-1.456282	0.1557
D(EXPCHIN(-1))	1.50E-10	8.89E-11	1.683592	0.1026
D(OFDI)	2.83E-07	5.12E-06	0.055276	0.9563
<b>D(EXR)</b>	<b>-0.016393</b>	<b>0.006865</b>	<b>-2.387858</b>	<b>0.0234</b>
C	-0.175573	0.068215	-2.573808	0.0152
R-squared	0.486519	Mean dependent var		-0.266650
Adjusted R-squared	0.298243	S.D. dependent var		0.362367
S.E. of regression	0.303559	Akaike info criterion		0.688472
Sum squared resid	2.764433	Schwarz criterion		1.184949
Log likelihood	-2.457906	Hannan-Quinn criter.		0.870450
F-statistic	2.584071	Durbin-Watson stat		1.718422
Prob(F-statistic)	0.019257			

Source: Compiled by authors.

Summing up our annual data analysis, the DOLS and ARDL estimations provide robust results which indicate that increases in the real effective yen exchange rate, i.e., appreciations, did have significant negative effects on the share of manufacturing in total employment in Japan. This is despite the fact that the yen also experiences longer periods of real effective depreciation and, thus, indicative of hysteresis effects (Krugman and Baldwin, 1987, Belke et al., 2013). DOLS estimations also find a negative and significant effect of outward FDI.

### 3.2 Analysis with monthly data, 1991-2017

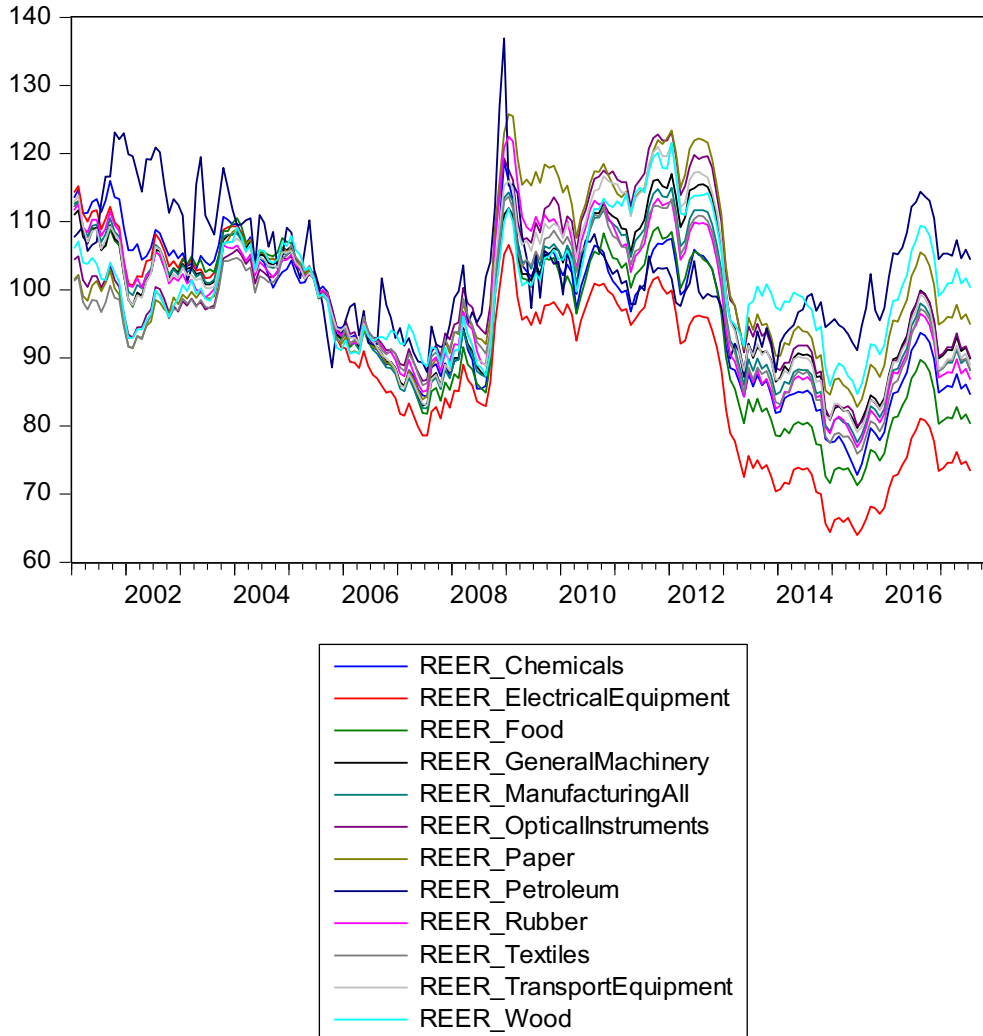
#### 3.2.1 Data and empirical model

In order to exploit our new data set on sector-specific yen exchange rates (Figure 6), we now estimate the real yen effective exchange rate effects on Japanese industrial production, using monthly data:

$$(2) \text{ INP}_X = \beta_0 + \beta_1 \text{ INP}_{X,t-1} + \beta_2 \text{ EX}_X + \beta_3 \text{ REER}_X + \beta_4 \text{ INDINPUTPRICE} + \beta_5 \text{ TFP} + \beta_6 \text{ INPUSA} + \varepsilon,$$

where  $\text{INP}_X$  = Industrial production by industry  $X$ ,  $\text{EX}_X$  = Exports by industry  $X$ ,  $\text{REER}_X$  = Industry-specific real effective exchange rate of industry  $X$  for 13 different industries,  $\text{INDINPUTPRICE}$  = Industrial input price, and  $\text{INPUSA}$  = Industrial production United States.<sup>7</sup> Appendix 2 provides details on the variables and sources.

**Figure 6: Real effective exchange rates for different industries**



*Source: Compiled by authors with data from RIETI (2017).*

Our data set covers the following sectors/industries: chemicals (CHEMICALS), electrical equipment (ELECTRICAL-EQUIPMENT), food (FOOD), general machinery (GENERAL-MACHINERY), optical instruments (OPTICAL-INSTRUMENTS), paper (PAPER), petroleum related products (PETROLEUM), rubber related products (RUBBER), textiles, (TEXTILES), transport equipment (TRANSPORTEQUIPMENT) and finally also wood related products (WOOD).

While all other variables are industry-specific,  $\text{INDINPUTPRICE}$  and  $\text{INPUSA}$  are non-industry-specific variables. In this sense, we follow a mixed panel-time series modelling approach. Figure 7 shows industrial production and industry-specific real effective exchange rates for selected industries. The

<sup>7</sup> The time series of industrial production have been seasonally adjusted, using the Census X-13 procedure.

two series tend to move in opposite direction, indicating that real effective exchange rates may indeed have a negative impact on industrial production.

**Figure 7: Industrial production and industry-specific real effective exchange rates for selected industries**



Source: Compiled by authors with data from RIETI (2017) and METI (2017).

### 3.2.2 Unit root tests

We now proceed with unit root tests of the variables listed below eq. (2) to be employed in our ARDL estimations and in our empirical mixed time series-panel models of Japanese industrial production which both necessitate the variables to be stationary.

The wide array of unit root tests we conducted, such as the ordinary Augmented-Dickey-Fuller test, the Dickey-Fuller GLS (ERS) test, the Phillips-Perron test, the Kwiatowski-Phillips-Schmidt-Shin test and the Ng-Perron unit root test, point into the direction of an I(1) process. Hence, we proceed with the individual variables considering them as I(1) variables which have to be differenced once in order to become stationary. The results are available on request.

We first conduct sector-specific Autoregressive Distributed Lag (ARDL) estimations based on first differences of the (non-pooled) variables.

### 3.2.3 Sector-specific autoregressive distributed lag-estimations for Japanese industrial production

The pattern of the results of our ARDL estimations of eq. (2) based on first differences of the (non-pooled) sector-specific and two further non-sector specific variables are displayed in Table 12. We rely on the same mechanics of ARDL models which have already been explained in section 3.1.4. This time, entries in the “selected model” column indicate the model selected according the Akaike information criterion. The first entry in brackets denote the lags of the *endogenous lagged variable*, the second entry corresponds with the lag number of the change of the *sector-specific exchange rate*, the third entry indicates the lag number of the change in *US industrial production* and the fourth one stands for the lag number of the change in Japanese *industrial input prices*. Deviating from the specification of eq. (2) we leave out sector-specific exports for reasons explained further below. Column 2 in Table 12 containing the yen exchange rate effect denotes the significant estimated coefficients for the *yen exchange rate* in the selected ARDL model. The *total* exchange rate effect (next column) is indicated in the third column of Table 12 by the *sum* of the period-specific significant entries. Values in brackets denote the respective significance level (probabilities). To illustrate, Table 13 contains a detailed ARDL estimation example, i.e. for the manufacturing sector. Its results are summarised in the first row of Table 12. Here, the impact coefficient of the exchange rate on the industrial production in the Japanese manufacturing sector is -0.61 with a rather high significance of lower than 0.01.

**Table 12: Impact of the real effective yen exchange rate on Japanese industrial production – A survey of estimated sector-specific ARDL models (excluding exports)**

Japanese industry sector	Selected ARDL model	Yen exchange rate effects	Yen exchange rate effect (total)
Manufacturing	ARDL(4, 3, 2, 0)	Lag 3: -0.61 (0.0032)	-0.61
Chemicals	ARDL(2, 3, 1, 0)	Lag 3: -0.41 (0.0036)	-0.41
Electrical equipment	ARDL(3, 3, 1, 2)	Lag 3: -1.04 (0.0101)	-1.04
Transport equipment	ARDL(3, 3, 2, 0)	Lag 3: -0.83 (0.0148)	-0.83
Rubber	ARDL(3, 3, 1, 0)	Lag 1: 0.44 (0.0981) Lag 3: -0.95 (0.0004)	-0.51
Optical instruments	ARDL(2, 3, 0, 0)	Lag 0: 0.90 (0.0294) Lag 3: 1.43 (0.0004)	-0.53
Paper	ARDL(4, 3, 1, 0)	Lag 3: -0.38 (0.0008)	-0.38
Food	ARDL(3, 0, 0, 0)	—	—
Textiles	ARDL(4, 3, 4, 0)	1: 0.27 (0.0785) 3: -0.25 (0.0830)	0.01

Source: Compiled by authors.

**Table 13: Impact of the real effective yen exchange rate on Japanese manufacturing industrial production – an estimated ARDL model**

Dependent Variable: D(INP\_MANUFACTURINGALL)  
 Method: ARDL  
 Sample (adjusted): 2001M06-2017M06  
 Included observations: 193 after adjustments  
 Maximum dependent lags: 4 (Automatic selection)  
 Model selection method: Akaike info criterion (AIC)  
 Dynamic regressors (4 lags, automatic): D(REER\_MANUFACTURINGALL) D(INPUSA) D(INDINPUTPRICE)  
 Fixed regressors: C  
 Number of models evaluated: 500  
 Selected Model: ARDL(4, 3, 2, 0)

Variable	Coefficient	Std. Error	t-Statistic	Prob.*
D(INP_MANUFACTURINGALL(-1))	-0.697699	0.072846	-9.577785	0.0000
D(INP_MANUFACTURINGALL(-2))	-0.867266	0.086039	-10.07993	0.0000
D(INP_MANUFACTURINGALL(-3))	-0.283453	0.087358	-3.244752	0.0014
D(INP_MANUFACTURINGALL(-4))	-0.165924	0.071672	-2.315035	0.0217
D(REER_MANUFACTURINGALL)	0.189543	0.219093	0.865127	0.3881
D(REER_MANUFACTURINGALL(-1))	0.213997	0.216040	0.990543	0.3232
D(REER_MANUFACTURINGALL(-2))	0.040311	0.213098	0.189164	0.8502
<b>D(REER_MANUFACTURINGALL(-3))</b>	<b>-0.614792</b>	<b>0.205586</b>	<b>-2.990434</b>	<b>0.0032</b>
D(INPUSA)	0.892882	0.661064	1.350675	0.1785
D(INPUSA(-1))	2.602084	0.653861	3.979568	0.0001
D(INPUSA(-2))	0.966521	0.681389	1.418456	0.1578
D(INDINPUTPRICE)	-0.013329	0.085517	-0.155863	0.8763
C	-0.227603	0.422910	-0.538183	0.5911
R-squared	0.549736	Mean dependent var	0.050777	
Adjusted R-squared	0.519718	S.D. dependent var	8.436238	
S.E. of regression	5.846512	Akaike info criterion	6.434549	
Sum squared resid	6152.707	Schwarz criterion	6.654316	
Log likelihood	-607.9340	Hannan-Quinn criter.	6.523548	
F-statistic	18.31378	Durbin-Watson stat	2.104685	
Prob(F-statistic)	0.000000			

Source: Compiled by authors.

The ordinary Pearson correlations conducted for the example of the Japanese manufacturing sector displayed in Table 14 demonstrate that *exports* of the Japanese manufacturing sector *dominate* any impact on the industrial *production of the same sector*. Our additional robustness checks show that we can *generalise* this statement for all sectors considered here: branch-specific exports dominate any impact on the respective sector's industrial production.

**Table 14: Ordinary Pearson correlations – the example of the Japanese manufacturing sector**

	D(INP_MANUFACTURINGALL)	D(EX_MANUFACTURINGALL)	D(REER_MANUFACTURINGALL)	D(INDINPUTPRICE)	D(INPUSA)
D(INP_MANUFACTURINGALL)	1.000000	0.826584	-0.008935	-0.058282	-0.011778

D(EX_MANUFACTURINGALL)	0.826584	1.000000	-0.103245	0.023029	0.092910
D(REER_MANUFACTURINGALL)	-0.008935	-0.103245	1.000000	-0.308087	-0.120676
D(INDINPUTPRICE)	-0.058282	0.023029	-0.308087	1.000000	0.241691
D(INPUSA)	-0.011778	0.092910	-0.120676	0.241691	1.000000

Source: Compiled by authors.

As a robustness check, we estimated sector-specific ARDL models, this time including Japanese exports. The results are surveyed in Table 15. The pattern of the results does not change materially.

**Table 15: Impact of the real effective yen exchange rate on Japanese industrial production – a survey of estimated sector-specific ARDL models (including exports)**

Japanese industry sector	Selected ARDL model	Yen exchange rate effects	Yen exchange rate effect (total)
Manufacturing	ARDL(4, 3, 4, 1, 0)	Lag 0: 0.49 (0.0001) Lag 1: 0.30 (0.0198) Lag 2: 0.26 (0.0377) Lag 4: -0.27 (0.0183)	0.78
Chemicals	ARDL(2, 0, 3, 0, 1)	Lag 3: -0.40 (0.0020)	-0.40
Electrical equipment	ARDL(3, 4, 4, 2, 4)	Lag 3: -0.74 (0.0180) Lag 4: -0.57 (0.0604)	-1.31
Transport equipment	ARDL(3, 2, 0, 0, 1)	Lag 0: 0.67 (0.0004)	0.67
Rubber	ARDL(4, 1, 3, 0, 1)	Lag 1: 0.53 (0.0268) Lag 3: -0.84 (0.0004)	-0.31
Optical instruments	ARDL(4, 6, 3, 0, 2)	Lag 0: 1.30 (0.0008) Lag 1: 0.80 (0.0382) Lag 3: -0.83 (0.0276)	1.27
Paper	ARDL(4, 4, 3, 3, 2)	Lag 3: -0.26 (0.0101)	-0.26
Food	ARDL(3, 4, 3, 0, 0)	Lag 3: -0.67	-0.67
Textiles	ARDL(4, 4, 0, 3, 3)	—	—

Source: Compiled by authors.

### 3.2.4 Panel estimations

As a first step, we test for unit roots in the pooled sector-specific time series. In order to make sure that we employ stationary variables in our panel estimations we conducted the following panel unit root tests for the pools of sector-specific variables: Levin, Lin and Chu (2002), Breitung (2000), Im, Pesaran and Shin (2003), Fisher-type tests using ADF and PP tests (Maddala and Wu 1999; Choi 2001), and Hadri (2000). All variables turned out to be I(1). The results are available on request.

As a second step, we conduct *Pooled Least Squares* estimations of a *mixed time series-cross section* model based on stationary time series with White cross-section standard errors (to allow for general contemporaneous correlation between the branch-specific residuals) and White covariance (White, 1980, MacKinnon and White, 1985). Non-zero covariances are allowed across cross-sections (degree-of-freedom corrected). The estimator we employ in this study is thus robust to cross-equation (contemporaneous) correlation between the branch-specific residuals and heteroskedasticity. The estimation results for our final model are displayed in Table 16.

**Table 16: Sectoral industrial production in Japan – determinants according to a Pooled Least Squares panel model**

Dependent Variable: D(INP?)  
 Method: Pooled Least Squares  
 Sample (adjusted): 2001M05-2017M05  
 Included observations: 193 after adjustments  
 Cross-sections included: 6  
 Total pool (balanced) observations: 1158  
 White cross-section standard errors & covariance (d.f. corrected)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.358571	0.337580	-1.062181	0.2884
D(INP?(-1))	-0.219148	0.047894	-4.575686	0.0000
D(EX?)	1.15E-08	1.30E-09	8.860615	0.0000
Sector-specific exchange rate coefficients:				
_CHEMICALS--D(REER_CHEMICALS(-3))	-0.360894	0.123608	-2.919660	0.0036
_OPTICALINSTRUMENTS-- D(REER_OPTICALINSTRUMENTS(-3))	-1.337488	0.315744	-4.235989	0.0000
_PAPER--D(REER_PAPER(-3))	-0.297193	0.138053	-2.152741	0.0315
_RUBBER--D(REER_RUBBER(-3))	-0.883009	0.247108	-3.573367	0.0004
_WOOD--D(REER_WOOD(-3))	-0.480935	0.178896	-2.688341	0.0073
_TEXTILES--D(REER_TEXTILES(-3))	-0.356062	0.127916	-2.783562	0.0055
Fixed Effects (Cross)	Not listed			
Effects Specification:				
Cross-section fixed (dummy variables)				
R-squared	0.214667	Mean dependent var	-0.221503	
Adjusted R-squared	0.205743	S.D. dependent var	7.903358	
S.E. of regression	7.043560	Akaike info criterion	6.754121	
Sum squared resid	56755.83	Schwarz criterion	6.815228	
Log likelihood	-3896.636	Hannan-Quinn criter.	6.777180	
F-statistic	24.05440	Durbin-Watson stat	2.173264	
Prob(F-statistic)	0.000000			

Source: Compiled by authors.

As a third step, we conduct *Pooled Estimated Generalised Least Squares* (EGLS, cross-section weights), again with White cross-section standard errors and White covariance. The results are summarised in Table 17. The results of our fixed effects redundancy test (at the beginning of Table 17) support our specific choice of fixed effects empirically.

**Table 17: Sectoral industrial production in Japan – determinants according to a Pooled EGLS panel model (with cross-section weights)**

Redundant Fixed Effects Tests  
 Test cross-section fixed effects





**Table 18: Sectoral industrial production in Japan – Determinants according to a Pooled EGLS panel model (cross-section SUR)**

Redundant Fixed Effects Tests				
Test cross-section fixed effects				
Effects Test		Statistic	d.f.	Prob.
Cross-section F		1.404752	(8,1717)	0.1895
Dependent Variable: D(INP?)				
Method: Pooled EGLS (Cross-section SUR)				
Sample (adjusted): 2001M05 2017M05				
Included observations: 193 after adjustments				
Cross-sections included: 9				
Total pool (balanced) observations: 1737				
Linear estimation after one-step weighting matrix				
White cross-section standard errors & covariance (d.f. corrected)				
Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.284434	0.201984	-1.408203	0.1593
D(INP?(-1))	-0.314740	0.021822	-14.42294	0.0000
D(EX?)	5.42E-09	1.92E-10	28.27803	0.0000
Sector-specific exchange rate coefficients:				
_CHEMICALS--D(REER_CHEMICALS(-3))	-0.299775	0.092382	-3.244931	0.0012
_ELECTRICALEQUIPMENT--				
D(REER_ELECTRICALEQUIPMENT(-3))	-0.493090	0.355530	-1.386914	0.1656
_GENERALMACHINERY--				
D(REER_GENERALMACHINERY(-3))	-0.340410	0.309889	-1.098489	0.2721
_OPTICALINSTRUMENTS--				
D(REER_OPTICALINSTRUMENTS(-3))	-1.304587	0.313259	-4.164558	0.0000
_PAPER--D(REER_PAPER(-3))	-0.209617	0.133158	-1.574201	0.1156
_RUBBER--D(REER_RUBBER(-3))	-0.754217	0.210364	-3.585299	0.0003
_TEXTILES--D(REER_TEXTILES(-3))	-0.281019	0.119584	-2.349969	0.0189
_TRANSPORTEQUIPMENT--				
D(REER_TRANSPORTEQUIPMENT(-3))	-0.387704	0.183868	-2.108603	0.0351
_WOOD--D(REER_WOOD(-3))	-0.376419	0.147706	-2.548435	0.0109
Fixed Effects (Cross)	Not listed			
Effects Specification				
Cross-section fixed (dummy variables)				
Weighted Statistics				
R-squared	0.361957	Mean dependent var	-0.029174	
Adjusted R-squared	0.354897	S.D. dependent var	1.216267	
S.E. of regression	0.976993	Sum squared resid	1638.903	
F-statistic	51.26540	Durbin-Watson stat	2.169783	
Prob(F-statistic)	0.000000			
Unweighted Statistics				
R-squared	0.437827	Mean dependent var	-0.142372	
Sum squared resid	114827.5	Durbin-Watson stat	2.336021	

Source: Compiled by authors.

Overall, the estimation results displayed in Tables 16 to 18 again appear well behaved with an eye on our main hypothesis. Significant negative sector-specific impacts of the yen effective real exchange rate can be established for at least five sectors (chemicals, optical instruments, rubber, textiles, wood), the exact number of sectors depending on the chosen estimation procedure.

According to our Pooled Least Squares estimations (Table 16) significantly negative effects of the (sector-specific) yen exchange rate on Japanese industrial production emerge for the chemicals, the optical instruments, the paper, the rubber, the wood and the textiles sector.

Our estimations based on Pooled EGLS (cross-section weights, Table 17) reveal that the (sectoral) yen exchange rate negatively impacts sectoral industrial production in the chemicals, the electrical equipment, the optical instruments, the paper, the rubber and the wood sector.

The largest number of sectors for which the yen exchange rate proves to be significant and reveals the expected negative sign emerges from our Pooled EGLS (Cross-section SUR) estimates (Table 18). Here, the expected negative sectoral exchange rate effect is identified for the chemicals, the electrical equipment, the general machinery, the optical instruments, the rubber, the textiles and the wood sector. We would like to stress that the broadest evidence of negative yen exchange rate impacts on Japanese sectoral industrial production is thus delivered by the most sophisticated estimation method applied by us in this paper: the feasible GLS specification correcting for both cross-section heteroskedasticity and contemporaneous correlation (cross-section SUR).

In all cases, the sectoral yen exchange rate enters the final model consistently with a *lag of three months* (time-to-build effect). What is more, the selected empirical models are rather *parsimonious* in terms of the number of variables included. The (changes in the) Japanese industrial input price (INDINPUTPRICE) and US Industrial production (INPUSA) are not part of the finally selected model. However, the lagged endogenous variable (change in Japanese industrial production) turns out to be highly significant throughout. Hence, in accordance with, for instance, Belke and Gros (2001), we feel legitimised to argue that our final empirical models can be interpreted as (at least partly) *reduced forms* against which we are able to separate the additional marginal impacts of the yen exchange rate and Japanese exports on industrial production.<sup>9</sup> Finally, critics may argue that the empirical realisations of the R-squared is not overly high. However, this is quite typical of regressions of changes on changes and the R-squared in our estimations appears to be located in the upper possible range.

#### *Summary of findings for monthly data analysis*

Both sector-specific ARDL estimations and estimations of a mixed time series-cross section model suggest small but highly significant effects of the REER on industrial production. In Section 3.2 we spared cointegration analysis because the time span 2000 to 2017 is much shorter than the one available in Section 3.1., i.e. 1970 to 2016.<sup>10</sup>

## **5. Conclusions**

In this paper we have investigated the role of the yen exchange rate in the de-industrialisation that has taken place in Japan over recent decades. There are certainly multiple factors that have played a role in this process, and we don't seek to evoke the impression that the exchange rate has been the single

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<sup>9</sup> This interpretation is valid for all finally selected model specifications in this paper which do not include the full set of variables proposed by theory, i.e., equations (1) and (2).

<sup>10</sup> Note that using quarterly data (Section 3.2) instead of annual ones (Section 3.1) does not help, because integration and cointegration are sample period properties and not necessarily dependent on the numbers of observation.

most important factor. Yet our results suggest that a strong yen had more than just transitory effects on Japanese manufacturing. The results also provide indication of hysteresis effects on manufacturing. Our findings are consistent with recent research on the hollowing out of the U.S. economy, where findings by Campbell (2017) also point to the presence of hysteresis effects. From a policy perspective, our findings suggest that central banks should try to avoid rapid, large-scale appreciations of the type Japan has seen a number of times over recent decades as negative effects cannot be easily reversed at a later stage. Yet Japan may also be special in the sense that its geographical proximity to emerging economies across East Asia facilitated outsourcing.

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## Appendix

### Appendix 1: Variables and sources for analysis with annual data, 1970-2016

Variable	Source
<i>EMPMAN</i> = Share of manufacturing in total number of employed persons (%)	Japan Labor Force Survey, Statistics Bureau
<i>CAP</i> = Gross domestic fixed capital formation (% of GDP)	World Bank World Development Indicators
<i>EXPMAN</i> = Manufacturing exports	World Bank World Development Indicators
<i>OFDI</i> = Outward direct investment	Japanese Trade and Investment Statistics, JETRO
<i>TFP</i> = Total factor productivity at constant national prices	Penn World Table 9.0
<i>EXPCHIN</i> = Chinese exports	World Bank World Development Indicators
<i>EXR</i> = Real effective exchange rate	Bank of Japan

### Appendix 2: Variables and sources for analysis with monthly data, 1991-2017

Variable	Source
<i>INP_X</i> = Industrial production by industry X	METI Survey of Production
<i>EX_X</i> = Exports by industry X	Global Trade Atlas
<i>REER_X</i> = Industry-specific real effective exchange rate of industry X for 13 different industries	Sato et al. (2013); RIETI, 2017
<i>INDINPUTPRICE</i> = Industrial input price	IMF Primary Commodity Prices
<i>INPUSA</i> = Industrial production United States	World Bank Global Economic Monitor