When, How Fast and by How much do Trade Costs change in the EURO Area?

Helmut Herwartz*
Henning Weber**

* Christian-Albrechts-Universität zu Kiel, Germany
**Freie Universität Berlin, Germany

BERLIN

ECONOMIC RISK

SFB 649

When, How Fast and by How Much do Trade Costs change in the Euro Area?

Helmut Herwartz[*] and Henning Weber[†]

October 30, 2008

Abstract

Microfoundations of the euro's effect on euro area trade hinge on the timing, the speed and the size of adjustment in trade costs. We estimate timing, speed and size of adjustment in trade costs for sectoral trade data. Our approach allows for sector specific impacts of trade costs on sectoral trade while controlling for unobserved but time-variant variables at the sector level. We find that, due to falling trade costs, trade within the euro area increases between the years 2000 and 2003 by 10 to 20 percent compared with trade between European countries that are not members of the euro area. Adjustment of individual sectors is extremely fast whereas aggregate adjustment spreads out because different sectors adjust at distinct times.

JEL Code: C31, C33, F13, F15, F33, F42
Keywords: Euro trade effect, gravity model, smooth transition, Kalman filter

1 Introduction

The literature that assesses the impact of the euro on euro area trade has converged to a fairly uniform empirical model. This conventional empirical model is based upon gravity theory which links equilibrium trade to trade costs, competitiveness and economic activity. Gravity theory is implemented as panel regression with trade as dependent variable and proxies for trade costs, competitiveness and economic activity as conditioning variables. The predominant purpose of the conventional model is to estimate by how much the level of euro area trade changes with the introduction of the euro. Commonly, this one time level shift in euro area trade is attributed to a corresponding one time level shift in trade costs. Accordingly, in the gravity regression proxies for trade costs comprise a euro dummy which indicates the euro introduction in January 1999.

However, empirical evidence on a level shift in euro area trade is insufficient to discriminate among competing theories on the euro's trade effect if all such theories happen to predict a reduction of trade costs and thus trade creation. In this case, additional characteristics of adjustment such as the timing and/or the speed by which trade adjusts to its new level are decisive to discriminate among competing theories. For instance, Baldwin (2006a) argues that the rapid increase in trade found in Micco, Stein, and Ordonez (2003) is difficult to reconcile with the construction of new plants to revert former foreign direct investment related to hedging. In turn, Baldwin and Taglioni (2005) argue that the increase in trade is rapid because reduction of exchange rate volatility induces a large number of small firms to enter export markets.

Our main contribution is to complement estimates of the 'By how much' with estimates of the 'When' and the 'How fast' of the change in euro area trade costs. We obtain such estimates by replacing the conventional level shift in trade costs with a smooth transition path. The smooth transition path allows the model to self-select the timing, the speed of adjustment and the size of the change in trade costs. In contrast, the conventional level shift, which remains a special case of our specification, allows the model to select the size of the change in trade costs but restricts this change to materialize instantaneously and exhaustively with the fixing of exchange rates in January 1999.

A priori, it seems difficult to justify the level shift in January 1999 in trade costs to which the conventional model resorts. Berger and Nitsch (2008b) argue that several major events are candidates for changes in euro area trade costs even though January 1999 has become the convention. For our sample from January 1995 to May 2006, candidate dates are the end of 1997 (the beginning of the third stage of the Economic and Monetary Union (EMU) is decided), January 1999 (factual start of third stage of EMU) and January 2002 (introduction of the euro as physical currency). Moreover, Micco, Stein, and

1

Ordonez (2003) and Flam and Nordström (2006b) report an increase in euro area trade already in 1998 and interpret this increase as anticipation effects. Smooth transition is flexible enough to handle a break point different from January 1999 because the model self-selects the timing of transition. DeNardis and Vicarelli (2007) argue that the euro may create trade effects in the medium and long run because home bias in preferences extends to former foreign markets in the course of time. In contrast to the instantaneous and exhaustive level shift in the conventional model, smooth transition has the potential to reflect delayed and gradual adjustment as well as long run effects because it does not force transition to complete within the estimation sample.

We undertake two more departures from the conventional model beyond smooth transition. Our second departure is motivated by the fact that the conventional model restricts the impact of trade costs to be homogeneous across countries. In particular, the conventional panel model comprises a cross section with countries both inside and outside the euro area. The euro dummy then is restricted to be identical for all euro area insiders and to be identical for all euro area outsiders. The difference between coefficients of the euro dummy for insiders and the euro dummy for outsiders is interpreted as the euro's trade effect. Parameter homogeneity also applies to trade cost variables others than the euro dummy. In this case, the impact of trade costs is restricted to be identical for the entire cross section.

We allow trade cost variables to have different impact across countries and, additionally, across trade sectors. To this end, we employ panel data that discriminate exports by partner country and by almost one hundred trade sectors. For instance, the data comprise German exports to France in the 'Iron and Steel' trade sector. Taking a mean group perspective (Pesaran and Smith (1995)), each trade sector in each trade relationship is allowed to respond differently to trade costs.

Anderson and van Wincoop (2004) expect effects of switching from national to common currency to differ considerably across countries. Suppose that one euro area country trades with countries both inside and outside the euro area whereas another euro area country trades mainly within the euro area. If switching to a common currency reduces trade costs within the euro area, the country with trade partners both inside and outside the euro area experiences a change in relative trade costs whereas the other country does not. Changes in relative trade costs redirect trade which constitutes a national euro effect.

Trade sector specific euro effects add to national effects. Taglioni (2002) emphasizes that the extent of vertical differentiation, the magnitude of economies of scale, the degree of industrial concentration, the size of non-tariff barriers and the relative location of reference markets and competitors differ substantially across sectors. Moreover, sector specific exposure to exchange rate risk as a result of pricing strategies or a size distribu-

tion of firms that differ across sectors potentially implies asymmetric euro effects.

Our last significant departure from the conventional model is due to the fact that the bulk of trade costs and competitiveness terms are either poorly measured or entirely unobserved. In the conventional model, time dummies common to the entire cross section serve as stand-in for all omitted or mismeasured variables. Baldwin (2006a) points out, however, that the use of common time dummies is at odds with gravity theory since measures of competitiveness are trade relationship specific according to this theory. At the sector level, gravity theory predicts sector specific competitiveness terms (Anderson and van Wincoop (2004)). Baldwin and Taglioni (2006) show that incomplete account of time variation in trade costs or competitiveness terms causes omitted variable bias in parameter estimates. Tractability appears a major reason to impose common time dummies because in the conventional model one set of time dummies for each trade relationship immediately exhausts all degrees of freedom.

We address omitted variable bias by treating time variant omitted trade costs and competitiveness terms as unobserved variables. Kalman filtering provides a conceptually straightforward account of this category of variables. In particular, the Kalman filter specification allows us to control for unobserved variation at the level of trade sectors and, at the same time, to address the short (data) history of the euro because the specification is very parsimonious. Admittedly, we exchange parsimony against some computational complexity since state space modeling requires numerical optimization.

Our main findings are as follows. Due to falling trade costs, trade within the euro area increases by 10 to 20 percent between the years 2000 and 2003 compared with trade between European countries that are not members of the euro area. Thus, contrary to the instantaneous and exhaustive level shift imposed by the conventional model, we find that it takes roughly three years for euro area trade to reach a new level. Moreover, euro area trade appears to start its transition roughly one year *after* the fixing of exchange rates in January 1999 so that our results do not conform easily with the convention to fix the break point in 1999. Moreover, our results neither confirm earlier findings of anticipatory activity nor suggest that there is much scope left for long run effects to develop.

Whereas adjustment of trade at the aggregate level takes about three years, adjustment of trade at the sector level is much faster. Aggregate adjustment is more spread out and gradual because different trade sectors adjust at different times. This conclusion holds true for trade within the euro area, trade among euro area outsiders and trade between the euro area and euro area outsiders. The high speed of adjustment at the sector level squares well with recent microfoundation of the euro's trade effect in Baldwin and Taglioni (2005). These authors argue that the reduction of exchange rate volatility induces considerable firm entry in export markets.

3

Baldwin (2006a) summarizes the literature on the trade effect of the euro as suggesting an increase in trade of 5 to 10 percent. Our mean estimates of the increase in euro area trade are roughly twice as large, even though confidence bands are consistent with an increase in trade of 5 to 10 percent. One explanation for the fact that we find larger effects is that the conventional euro dummy kicks in too early so that the conventional size estimate is biased downwards. Interestingly, the size of the increase in trade also falls when we restrict parameters to be homogeneous across trade sectors. However, when testing a model, which allows for parameter heterogeneity at the sector level, against a model with homogeneous parameters, we find strong evidence for sector specific coefficients. Disregarding sectoral heterogeneity thus leads to a downward bias in the extent by which trade costs are estimated to fall and, accordingly, by which the level of trade increases.

One interpretation which is consistent with the reduction in trade costs and, correspondingly, with the type of trade creation described here is that the euro led to lower trade costs thereby creating euro area trade in the course of time. Baldwin (2006a) contemplates both systematic measurement error in trade statistics and the devaluation of the euro in its early days as alternative interpretations consistent with trade creation around the time of the euro introduction. Systematic measurement error in trade statistics can arise because VAT taxes induce incentives to overstate export and understate import figures. To check for the impact of fraud, we contrast results obtained for export data with results obtained when using import data. We find fairly similar results in the two cases, both qualitatively and quantitatively, even though import data overall appears less informative. At least the more simple fraud schemes thus do not appear major drivers of our results. Alternatively, the depreciation of the euro right after its introduction may have induced euro area expenditure to shift from foreign products to euro area products thereby constituting euro area trade creation. In our estimates, we control for this possibility and include several measures of exchange rates into the set of conditioning variables.

Several papers in the literature on the trade effect of the euro, which is comprehensively reviewed in Baldwin (2006a) and Baldwin (2006b), directly relate to our work. Micco, Stein, and Ordonez (2003), Flam and Nordström (2006a), Flam and Nordström (2006b) and DeNardis and Vicarelli (2007) make inference on the timing of the euro's trade effect by using repeated year dummies interacted with a euro dummy. Baldwin, Skudelny, and Taglioni (2005), Flam and Nordström (2006b) and Nitsch and Pisu (2008) work with sectoral data and estimate sector specific and in some cases also trade relationship specific coefficients. From perspective of gravity theory, one drawback of these studies is that they either resort to common time dummies to control for time variant omitted variables or do not control for time variant omitted variables at all.

2 Theory

Gravity theory (Anderson and van Wincoop (2003), Anderson and van Wincoop (2004), Baldwin (2006a)) relates equilibrium exports to the product of foreign expenditure and home production and to trade costs relative to competitiveness terms,

$$X_k^{ij} = E_k^j Y_k^i \left(\frac{\tau_k^{ij}}{\Pi_k^i P_k^j} \right)^{1-\sigma_k}. \tag{1}$$

Nominal equilibrium exports of reporter country i to partner country j in sector k are denoted X_k^{ij}, nominal expenditure in this sector is E_k^j, and nominal production is Y_k^i. The sectoral elasticity of substitution is $\sigma_k > 1$. The trade cost function τ_k^{ij} summarizes all trade costs and is specified below. If bilateral trade costs τ_k^{ij} are reduced say because a common currency decreases transaction costs exports from i to j increase. The variables P_k^j and Π_k^i represent inward and outward competitiveness terms, respectively. Both terms are weighted averages of bilateral trade costs relative to the welfare based price levels of the respective trading partner. Weights reflect the size of a sector. In particular, if importing country j faces high trade costs with respect to exporters other than i this increases inward competitiveness P_k^j and exports from i to j increase. Outward competitiveness Π_k^i reflects the notion that, if from i's perspective trade costs are higher for markets other than j, more will be exported from i to j.

Neither competitiveness terms nor trade costs have easy to access empirical correspondents. For instance, gravity consistent competitiveness terms require a stand on which foreign products compete with national products. Since we work with a small trade matrix comprising Germany, France, Italy, United Kingdom, Sweden and Denmark we are likely to miss many important substitutes. Moreover, Baier and Bergstrand (2001), Feenstra (2003) and Anderson and van Wincoop (2004) point out that measured export and import price indices do not align with competitiveness terms because, for instance, such indices do not reflect home bias in consumer preferences. Adding to these difficulties, we are likely to omit many important trade cost variables since detailed trade cost data is rare in general and even harder to obtain for a panel of trade flows at annual or even monthly frequency and at sector level. Anderson and van Wincoop (2004) survey trade costs and their availability.

3 Empirical Model

3.1 Basic Setup

We interpret gravity theory in equation (1) as state space system. This interpretation provides a conceptually straightforward account of unobserved trade costs and competitiveness terms. The log linear equation (1) jointly with a specification of measurable trade costs represents the observation equation of the state space system. Competitiveness terms and omitted trade cost variables are absorbed into the state equation. In this section we describe the most general panel model we take into consideration. When conducting specification tests below we describe the restrictions imposed to arrive at less general models.

Let i denote reporters, j partners, k trade sectors and t time. Subsume reporter and partner indices under the trade relationship index $s = ij, i \neq j$ and let S (or K or T) denote the maximum number of trade relationships (or sectors or observations). For each trade relationship s we estimate the following model,

$$y_{kt}^{(s)} = q_{it}\beta_{ik}^{(s)} + q_{jt}\beta_{jk}^{(s)} + (1 - \sigma^{(s)})[\ln(\tau_{kt}^{(s)}) + \lambda_{kt}^{(s)}] + u_{kt}^{(s)}, \tag{2}$$

$$\ln(\tau_t^{(s)}) = \theta_{0k}^{(s)}\left[1 + \exp\{-\theta_{1k}^{(s)}(\tilde{t} - \zeta_k^{(s)})\}\right]^{-1} + (Z_t^{(s)})'\gamma_k^{(s)} + c^{(s)}, \tag{3}$$

$$\lambda_{kt}^{(s)} = \lambda_{kt-1}^{(s)} + v_{kt}^{(s)}, \tag{4}$$

$$u_{kt}^{(s)} \sim N(0, g_k^{(s)}) \quad, \quad v_{kt}^{(s)} \sim N(0, h_k^{(s)}) \quad, \quad E[u_{kt}^{(s)}v_{kr}^{(s)}] = 0 \quad \forall t, r. \tag{5}$$

The observation equation (2) specifies the log of sector k exports $y_{kt}^{(s)}$ for trade relationship s conditional on scale variables q_{it} and q_{jt}, the log of measurable trade costs $\tau_{kt}^{(s)}$ and the log of unobserved trade costs and competitiveness terms $\lambda_{kt}^{(s)}$.

Equation (3) formalizes the log of measurable trade costs as a smooth transition path, contained in square brackets, plus measurable control variables $Z_t^{(s)}$ and a constant term $c^{(s)}$. Following Luukkonen, Saikkonen, and Teräsvirta (1988), we model smooth transition as a logistic distribution function where $\theta_{0k}^{(s)}$ measures the size of transition while $\theta_{1k}^{(s)} > 0$ governs the speed of transition. In order to immunize $\theta_{1k}^{(s)}$ against the scale of the time index t the latter is standardized as $\tilde{t} = t/(T \times \sqrt{0.08333})$ (Bauwens, Lubrano, and Richard (2000)). The transition path is centered around the coefficient $\zeta_k^{(s)}$ which governs the timing of transition.

Jointly, estimates of the three coefficients $\theta_{0k}^{(s)}, \theta_{1k}^{(s)}$ and $\zeta_k^{(s)}$ completely describe transition. For instance, a large positive estimate of $\theta_{0k}^{(s)}$ implies a large size of transition; a large positive estimate of $\theta_{1k}^{(s)}$ implies fast adjustment; a large positive estimate of $\zeta_k^{(s)}$ implies transition that takes place late in our estimation sample. The smooth transition path comprises an instantaneous and exhaustive level shift in trade costs as special case.

The state equation (4) specifies the evolution of sector specific unobserved variables as random walk. The random walk assumption implies that unobserved variables may exhibit stochastic trends over time. Mongelli, Dorrucci, and Agur (2005) and Berger and Nitsch (2008b) argue that European integration is continuously deepening according to indices on institutional integration, trade liberalization and tariffs. To the extent that this integration process is not captured by measurable trade costs, unobserved trade costs are likely to reflect ongoing European integration in our model. Berger and Nitsch (2008b) approximate the integration process with a deterministic time trend common to all trade relationships s. However, our stochastic trend assumption specific to trade sectors appears to better describe this process. We assume that residuals of the observation equation $u_{kt}^{(s)}$ and of the state equation $v_{kt}^{(s)}$ are uncorrelated at any lead or lag.

Control variables $Z_t^{(s)}$ include real effective exchange rates of the reporting country $reer_t^{(s)}$, real bilateral exchange rates between both trading partners $rex_t^{(s)}$ and real exchange rates of the reporting country relative to the U.S. $rexus_t^{(s)}$. Adding bilateral exchange rates among trading partners and with the U.S. separately gives a prominent role to variation in these two prices beyond their appearance in the real effective exchange rate. Appendix D describes the construction of these variables. Flam and Nordström (2006b) and Baldwin (2006a) emphasize that exchange rates help to discriminate potential expenditure switching effects due to changes in international relative prices from effects of introducing the common currency. When the euro depreciated after its introduction products sold in euro became cheaper relative to products sold e.g. in U.S. dollar. Potentially, part of euro area demand for foreign products was redirected back to the euro area due to this change in international relative prices. Thus, if not controlled for, the model may falsely attribute the change in euro area trade due to a change in international relative prices to the introduction of the common currency. Also, $Z_t^{(s)}$ includes a measure of exchange rate volatility $vol_t^{(s)}$ to control for a potential link between exchange rate risk and trade. We describe construction of exchange rate volatility below. Finally, we add an index of energy prices $en_t^{(s)}$ as a measurable proxy for transportation costs.

The dependent variable $y_{kt}^{(s)}$ in equation (2) is likely to be nonstationary according to frequently inferred time series features.[1] In this case the empirical model may suffer from

---

[1]Unit root tests powerfully underscore the likelihood of stochastic trends in $y_{kt}^{(s)}$. In the light of the

spurious findings in the sense that coefficient estimates of nonstationary right hand side variables fail consistency. Accordingly, balancing the regression model (2) – (5) requires at least one nonstationary variable on the right hand side. We regard scale variables q_{it} and q_{jt} and exchange rate measures as candidates to cointegrate with the dependent variable. Moreover, the unobserved state variable in (4) evolves nonstationary and is a further candidate for cointegration.

To guard against spurious regression we diagnose the stochastic features of model implied residuals $u_{kt}^{(s)}$. In case the latter residual processes are stationary, variables entering the observation equation are either stationary or nonstationary but cointegrated processes. Chang, Miller, and Park (2008) derive that the common (Quasi) Maximum Likelihood (QML) interpretation of modeling stationary processes by means of the Kalman filter also applies for multivariate nonstationary processes sharing a common trend. As a consequence the validity of standard specification tests as e.g. likelihood ratio (LR) tests does not rely on the stationarity of $y_{kt}^{(s)}$ or conditioning variables.

3.2 Functional Coefficients and Estimation

We collect all coefficients in the two vectors

$$\boldsymbol{\psi}_k^{(s)} = \left(\beta_{ik}^{(s)}, \beta_{jk}^{(s)}, \theta_{0k}^{(s)}, \theta_{1k}^{(s)}, \zeta_k^{(s)}, \boldsymbol{\gamma}_k^{(s)\prime}, h_k^{(s)}, g_k^{(s)} \right)' \quad \text{and} \quad \boldsymbol{\phi}^{(s)} = (\sigma^{(s)}, c^{(s)})',$$

where $\boldsymbol{\psi}_k^{(s)}$ comprises sector specific coefficients and $\boldsymbol{\phi}^{(s)}$ comprises coefficients not specific to sectors. To estimate sector specific coefficients $\boldsymbol{\psi}_k^{(s)}$ we presume a parsimonious functional representation in which sector specific coefficients equal a common intercept term and slope coefficient multiplied by a sector specific scalar $a_k^{(s)}$,

$$\boldsymbol{\psi}_k^{(s)} = (\mathbf{1} + \boldsymbol{\psi}_1^{(s)} a_k^{(s)}) \odot \boldsymbol{\psi}_0^{(s)} . \tag{6}$$

Here $\mathbf{1}$ is a unit vector of appropriate dimension and $\boldsymbol{\psi}_1^{(s)}$ and $\boldsymbol{\psi}_0^{(s)}$ are vectors of unconditional coefficients. The operator $'\odot'$ signifies 'element-by-element' multiplication and the scalar $a_k^{(s)}$ with $\sum_k a_k^{(s)} = 1$ reflects the importance of sector k in reporting country i. To be precise, denote the relative average quantity traded in sector k as weight $w_k^{(s)}$ and denote the rank associated with $w_k^{(s)}$ conditional on s with $\tilde{w}_k^{(s)}$. The importance of sector k then is $\tilde{a}_k^{(s)} = \tilde{w}_k^{(s)} / (\sum_k \tilde{w}_k^{(s)})$ and a corresponding mean zero weighting sequence is $a_k^{(s)} = \tilde{a}_k^{(s)} - (1/K) \sum_k \tilde{a}_k^{(s)}$. Appendix A describes computation of weights.

---

plentitude of time series entering the empirical models we refrain from providing detailed results on unit root testing. Doing so reveals that almost uniformly first differences of employed time series are stationary so that the highest order of stochastic trending is one.

As alternative to the rank-based functional variable we also experimented with replacing $a_k^{(s)}$ by original weights $w_k^{(s)}$. As a purely descriptive matter, however, the distribution of weights is heavily skewed. There are only a few very large sectors but many small ones. When replacing $a_k^{(s)}$ by $w_k^{(s)}$ this skewness is amplified by the fact that sector specific coefficients and thereby functional variables enter the model in a nonlinear fashion. Therefore, functional estimates were dominated by the few very large sectors almost uniformly over all trade relationships.

Equation (2) restricts the elasticity of substitution $\sigma^{(s)}$ to remain common for all sectors conditional on trade relationship s. A linear functional relationship in the elasticity of substitution instead creates a quadratic functional relationship in export elasticities with respect to e.g. control variables contained in $Z_t^{(s)}$. The quadratic functional relationship follows because export elasticities comprise the product of $\sigma^{(s)}$ with coefficients which already depend linearly on the functional variable $a_k^{(s)}$. We avoid such quadratic relationships by imposing $\sigma_k^{(s)} = \sigma^{(s)}$.

Ideally, sector specific coefficients $\boldsymbol{\psi}_k^{(s)}$ should be flexible enough to reflect the many dimensions which make sectors transit differently and respond differently to changes in control variables. To name only a few, sectors are likely to differ with respect to the tightness of competition, with respect to pricing strategies and strategic interaction or regarding the degree of product substitutability. Our functional specification pretends that all relevant dimensions are reasonably well represented by a sector's market size which obviously is not the case. However, besides being an operative measure we believe that market size is a useful functional variable in that it correlates with at least the more relevant dimensions.[2]

QML estimation of the empirical model deserves iterative optimization due to non-linearities in model parameters and the presence of the unobserved processes $\lambda_{kt}^{(s)}$. A few parameters entering the state space model are estimated conditional on a restricted support. First, variance parameters are determined as exponentials of underlying parameters to ensure positive variation measures,

$$g_k^{(s)} = \exp\left(\underline{g_k^{(s)}}\right) \quad \text{and} \quad h^{(s)} = \exp\left(\underline{h_k^{(s)}}\right),$$

where underlining signifies that log likelihood optimization is done, for instance, with respect to $\underline{g_k^{(s)}}$ rather than $g_k^{(s)}$. The second set of restrictions applies to coefficients of the smooth transition function. One observes that the term in squared brackets in equation

---

[2]Alternative functional variables would be number of firms, profits, markups, exchange rate pass through or fixed costs of production. Besides data availability, relying on several functional variables would considerably boost the parameter space and render optimization a rather challenging task. Factor analysis is one means to reduce dimension and may turn out an interesting extension of the setup considered here.

(3) degenerates to a constant as $\theta_{k1}^{(s)}$ tends to zero. The state space model might lack identification in this case because trade costs already comprise a constant term. Accordingly, we restrict the support of $\theta_{k1}^{(s)}$ to strictly positive values. Fastest 90% of transition is restricted to one month by imposing $\theta_{k1}^{(s)}$ smaller than 232.89. Also, we impose bounds on the parameter space of the timing parameter $\zeta_k^{(s)}$ to prevent the transition function from isolating the first twelve and the last twelve sample observations. Within these bounds the timing parameter is free to adjust. We implement parameter restrictions with the cumulative Gaussian Φ, $0 < \Phi < 1$, as

$$\theta_{1k}^{(s)} = 232.89\ \Phi(\underline{\theta_{1k}^{(s)}}) \quad \text{and} \quad \zeta_k^{(s)} = 0.30343 + 2.8826\ \Phi(\underline{\zeta_k^{(s)}}).$$

Appendix A provides details to obtain such bounds. Finally, we ensure $\sigma^{(s)} > 1$ in line with economic theory. For optimizing over explicit or underlying parameters obtaining $(\boldsymbol{\psi}_k^{(s)\prime}, \boldsymbol{\phi}^{(s)\prime})\prime$ the *optmum* routine in GAUSS is used.

4   Data

We use monthly bilateral export data from January 1995 to May 2006 (137 months). We rely on monthly data to collect as much information as possible around the hypothesized break point. In EUROSTAT's COMEXT database export data is available in value (current euro) and volume (tons) and is disaggregated according to the HS two digit level. The HS classification provides a break down of aggregate trade into 99 trade sectors of which we consider $K = 96$.[3] For estimation we convert export and import data into year 2000 euros.

Baldwin (2006a) argues that trade figures are systematically distorted because VAT taxes induce misreporting incentives, and that this distortion may be particularly relevant around the time of the euro introduction. Berger and Nitsch (2008a) show that the gap between export and import statistics correlates positively with measures of corruption. Typically, misreporting due to VAT induces trade statistics to overstate export figures and to understate import figures though some fraud schemes may have considerably more complex implications. A first check for the impact of fraud is to contrast results obtained with export data to results obtained when using import data. Therefore, we explore our specification of gravity for both export and import data. Import data is also drawn from COMEXT.

---

[3]See http://fd.comext.eurostat.cec.eu.int/xtweb/ for COMEXT database. Sectors 77 and 98 do not contain any data for our sample. Also we drop sector 99 ('Other Products'). HS abbreviates Harmonized Commodity Description and Coding System. Baldwin and Taglioni (2006) address a wide range of possible misspecifications of the gravity equation related to data construction.

At the two digit level trade data for some sectors is plagued by irregularly missing observations. We do not exclude such sectors from our analysis since Kalman filter recursions are easily modified to cope with irregularly missing observations. Appendix B provides details on the employed Kalman filter. Hence, the empirical analysis does not suffer from imputed measures replacing missing observations and is not subject to sample selection bias as a consequence of excluding a potentially nonrandom fraction of trade sectors from the analysis.

Our trade matrix comprises Germany, France, Italy, United Kingdom, Sweden and Denmark so that we obtain $S = 30$ trade relationships. We restrict attention to European Union (EU) member states to maintain as much homogeneity as possible along the country dimension. EU countries have been subjected to similar legislation and regulation in the wake of the European Single Market initiative after 1993. Germany, France and Italy cover a major fraction of the euro area both in terms of population and in terms of GDP. United Kingdom, Sweden and Denmark constitute a useful reference group because they largely share the European history except for the most recent episode of the common European currency.

Out of the 30 trade relationships six involve countries which both have adopted the euro (U2), six involve countries which both have not adopted the euro (O2), nine involve countries where the reporting country has adopted the euro but the partner country has national currency (OUT) and nine involve countries where the partner country has adopted the euro but the reporter has not (IN). We compute mean group estimates based on these subsets. The first column of table 1 lists all trade relationships explicitly.

Gravity theory suggests to use data on sector production and sector expenditure for scale variables. We use indices of industrial production as proxy for sector production and sector expenditure but allow for sector specific coefficients $\beta_{ik}^{(s)}$ and $\beta_{jk}^{(s)}$ to mitigate inferior data quality. Baldwin, Skudelny, and Taglioni (2005) estimate gravity equations for two digit trade sectors and compare a specification with sector specific data on gross value added with a specification which uses aggregate GDP data as proxy for sectoral activity allowing for sector specific coefficients. They find that conclusions about the size of the change in trade costs are sensitive with respect to the proxy for sectoral activity but point to difficulties to obtain disaggregated data for gross value added. Flam and Nordström (2006b) report similar data problems with one digit annual data and estimate sector specific regressions with aggregate GDP data. In our case data availability is even more a constraint due to the monthly data frequency which motivates the choice of industrial production as proxy for sectoral activity.

We measure nominal exchange rate volatility nonparametrically as

$$
\left(vol_t^{(s)}\right)^2 = \frac{1}{D_t} \sum_{d=1}^{D_t} \left(\Delta \ln e_d^{(s)} - \frac{1}{D_t} \sum_{d=1}^{D_t} \Delta \ln e_d^{(s)}\right)^2 . \tag{7}
$$

Here $e_d^{(s)}$ represents daily quotes of reporter i's currency in terms of partner j's currency and D_t is the number of days per month. A similar measure based on weekly data is proposed in Baldwin, Skudelny, and Taglioni (2005). Appendix D provides further details on data construction and sources.

5 Model Selection and Diagnostic Checking

In this section, we impose and test particular restrictions on the functional smooth transition state space model outlined in section 3. First, we test the smooth transition path against a euro dummy. Second, we consider the exclusion of unobserved variables. Both restrictions are imposed in models with coefficients that are homogeneous across sectors and without exogenous control variables. Third, the homogeneous smooth transition state space model (including the smooth transition path and allowing for unobserved variables) is contrasted against the functional smooth transition state space model which allows coefficients to differ across sectors. The functional specification turns out preferable. Fourth, we do extensive residual checking for the functional smooth transition state space model. Finally, we illustrate the explanatory content of control variables. After this sequence of tests, we chose the functional smooth transition state space model including exogenous control variables as our preferred specification.

When it comes to diagnostic checking, a particular modeling issue is to guard against the potential of spurious regressions because stochastic trends are likely to govern the dependent as well as explanatory variables in equation (2). For this purpose we test the null hypothesis of nonstationarity for estimated residuals of the observation equation. Residual based testing for unit roots compares standard ADF statistics with a 5% critical value of -4.74 (Fuller (1976)) which is the relevant critical value for testing residuals of static regressions involving 6 nonstationary variables. The lag order of the ADF regression is 3 throughout. The critical value is likely conservative for the considered testing problem because the number of potential nonstationary right hand side variables in the state space model is 3 when excluding exogenous control variables but including unobserved variables. Therefore and because estimated error sequences subjected to testing are not obtained from static cointegrating regressions, ADF tests provide a rather descriptive view at overall model reliability. For each trade relationship, the analysis covers dynamics for

96 trade sectors. Therefore, we document empirical frequencies of rejections of the unit root null hypothesis rather than unit root statistics at sectoral level.[4]

In a similar vein as outlined for unit root testing we document diagnostic results to evaluate if model residuals feature serial correlation. Testing against serial correlation by means of a heteroskedasticity robust Wald statistic is detailed in appendix C.

Results from model selection and diagnostic checking are documented in table 1. By line, the table displays results for particular trade relationships. For model selection, we mostly employ likelihood ratio (LR) tests. The fact that we consider a set of 30 trade relationships for both imports and exports adds complexity to the provision of empirical results. In light of the plentitude of estimated empirical models, space considerations only allow a structured and condensed overview of particular model features. We now discuss the particular issues regarding model selection and diagnostic checking in detail.

(i) Smooth Transition: To assess the marginal contribution of the flexible smooth transition path in comparison with a conventional level shift in trade costs, the state space model with homogeneous coefficients ($\boldsymbol{\psi}_1^{(s)} = 0$) is alternatively estimated under restrictions $\theta_{1k}^{(s)} = 232.89$ and $\zeta_k^{(s)} = 1.2137$. These restrictions closely approximate a level shift in January 1999. For numerous trade relationships the smooth transition specification is supported by LR_d statistics in table 1 significant at conventional levels. Out of 30 LR_d statistics, 10 and 12 (5 and 5) are significant at the 10% (5%) level when modeling imports and exports, respectively. LR tests inferring against the euro dummy specification are interpreted to follow an asymptotic χ^2 distribution. However, the true distribution of LR_d statistics is actually unknown because the homogeneous state space model itself is likely to be too restrictive. Consequently, these specification tests should be treated with care and are rather a descriptive measure of model accuracy.

(ii) Unobserved Variables: To describe the explanatory content of unobserved variables $\lambda_{kt}^{(s)}$ we estimate smooth transition regression models consisting of equations (2) and (3) but excluding the state equation (4). That is, we impose $\lambda_{kt}^{(s)} = 0, \forall s, k, t$. We then compare the resulting standard error estimate $\sqrt{\hat{g}^{(s)}}$ with the corresponding quantity obtained from the homogeneous state space model. In table 1, SR denotes the ratio of these standard error estimates. Although this measure is purely descriptive it strongly underpins

---

[4]Tests on stationarity are only performed for sectors where the number of missing observations is at most 5. Consequently the frequencies of rejections of H_0 documented in table 1 refer to populations that depend on the trade relationship. In fact the number of diagnostic tests varies between 48 and 96 (53 and 96) for the analysis of import (export) relationships. The minimum numbers of 'complete' trade time series are obtained when modeling Italian imports from Sweden or Swedish exports to Italy.

the explanatory content of unobserved variables. Excluding sector specific unobserved variables involves a magnification of implied error variations by factors of 7.3 (imports of the United Kingdom from Denmark) up to 56.25 (French imports from Germany or French exports to Germany). At the first sight these factors appear unreasonably large. However, it is intuitive to regard the sector specific unobserved variables to potentially cointegrate with the sector specific trade variables. Therefore, a model without unobserved variables is likely to yield nonstationary residuals, so that inferential conclusions might be spurious. Accordingly, the marked reduction of variance estimates achieved by means of the state space representation becomes plausible.

(iii) Sector Specific Coefficients: Most strikingly, the functional state space model with sector specific coefficients is uniformly and significantly supported when compared to the homogeneous state space model. Introducing 7 additional parameters, smallest LR_f statistics testing the functional against the homogeneous state space model ($H_0 : \boldsymbol{\psi}_1^{(s)} = 0$ versus $H_1 : \boldsymbol{\psi}_1^{(s)} \neq 0$) are 635.7 and 734.6 for export and import data, respectively. These statistics could be compared with critical values from a $\chi^2(7)$ distribution. However, all statistics are in favor of the functional state space model at any conventional significance level.

(iv) Stochastic Trends and Serial Correlation: The likelihood of stochastic trends featuring model implied residuals is rather limited for both import or export data. The $I(1)$ columns in table 1 indicate empirical frequencies of rejections of the unit root null hypothesis. For most trade relationships in the functional state space model, almost all estimated sector specific residual sequences are found stationary. Thus, the conditional model is successful in filtering out common stochastic trends so that spurious regression results are unlikely. The evidence against remaining stochastic trends is similarly weak when modeling trade dynamics by means of homogeneous conditional models. However, we do not report model diagnostics for the homogeneous state space model for reasons of space.

First order residual correlation is detected for about one third of sector specific residual series for the functional state space model.[5] This is evident from AR1 columns in

---

[5]Testing against joint autocorrelation at lag 1 to 12, the empirical evidence against serially uncorrelated model residuals is even stronger. When comparing the functional against the homogeneous state space model (not reported) it is noteworthy that diagnostic model features also support the more general model class. Almost uniformly the frequency of significant autocorrelation test statistics is lower for the functional as it is for the homogeneous state space model. For the special case of Danish imports from Germany we find that almost all (86 out of 93) sector specific residual sequences feature first order autocorrelation. We treat this diagnosis as a hint at potential model misspecification or computational obstacles and remove

14

table 1 which indicate empirical frequencies of rejections of the null hypothesis of no serial correlation. QML parameter estimates are inefficient but remain unbiased in case of serially correlated error terms since lagged dependent variables are not included as explanatory variables. However, to us the efficiency aspect is of minor importance because the interpretation of estimation results relies on mean group estimators obtained as weighted averages of ML estimates (Pesaran and Smith (1995)). For the particular schemes employed to determine weighted mean group estimates see appendix A.

(v) Exogenous Control Variables: Augmenting the functional model jointly with all additional control variables $vol_t^{(s)}, reer_t^{(s)}, rex_t^{(s)}, rexus_t^{(s)}$ and $en_t^{(s)}$ assuming $\boldsymbol{\gamma}_k^{(s)} \neq 0$ delivers LR_X statistics in table 1 that are mostly significant. At a 5% significance level, 21 and 18 (out of 30) trade relationships are improved by including additional control variables. Marginal contributions of individual control variables are discussed when it comes to the economic discussion of the mean group parameter estimates in the next section. We also assess the effect of each control variable separately on the transition path for the homogeneous and the functional state space model. These results (not provided here) indicate that in particular the timing of transition remains a very robust feature of our estimates.

6 Results

We discuss results of our preferred model which is the functional smooth transition state space model covering the full set of exogenous control variables. First, we inspect mean group coefficient estimates others than those determining transition. Then we turn to the characteristics of the transition path.

6.1 Coefficient Estimates

Table 2 reports estimated coefficients for our preferred model specification for both export and import data. The column labeled ALL provides mean group coefficients. Mean group coefficients are weighted average coefficients over all trade sectors and over all trade relationships where weights correspond to the relative sector size. Subsequent columns provide mean group coefficients averaged over the various subsets of trade relationships U2, O2, IN and OUT. Subsets are described in section 4. Appendix A provides details on the computation of mean group estimates and corresponding standard errors.

---

this particular trade relationship from the sample when it comes to discussing the material implications of the estimated models in section 6.

Estimates of the elasticity of substitution σ are around 5 for both export and import data and across the various subsets of trade relationships.[6] Broda and Weinstein (2004) report elasticities of substitution around 4 using SITC three digit U.S. data for the period 1990–2001. Our estimates around 5 refer to two digit European data for a slightly different classification (HS instead of SITC) but overall appear to comply with estimates in Broda and Weinstein (2004). This similarity is reassuring in light of the fact that our state space setup does not make use of data on international prices.

In line with theory industrial production in reporter (β_i) and partner (β_j) country significantly increases exports (imports) which holds true over all trade relationships on average and for the majority of subsets. Though theory predicts a unit elasticity, estimates which differ from unity may signal ongoing change in the ratio of sectoral exports (imports) over industrial production. This explanation pairs well with the high estimation accuracy for some coefficients. Also, Baldwin (2006a) argues that inferior data quality may be one factor pushing elasticities below unity. Indeed, our proxy of sectoral activity is identical for all sectors. Even though industrial production certainly is a reasonably accurate measure for economic activity in some sectors (say, 'Manufacturing') it is likely to be less so for others (say, 'Food').

Exchange rate volatility is usually considered an impediment to trade and in particular so for small firms without the financial statue to hedge exchange rate risk.[7] In our sample, exchange rate volatility *vol* has a tendency to foster trade due to the small positive elasticity of 0.014, even though the positive relationship is not a robustly significant feature of the estimates. Theories which draw on the option value of trade predict such a positive link between trade and exchange rate uncertainty. However, here it appears more likely that the tendency of coefficient estimates to become positive relates the trend decrease in exchange rate volatility due to the convergence process towards the third stage of EMU to a slight moderation in euro area trade in the first half of our sample period.

A high effective exchange rate *reer* decreases exports with elasticity -0.2 when taken over all trade relationships. If domestic goods become more expensive relative to a weighted basket of foreign goods this reduces exports. However, the evidence over subsets is mixed so that in the aggregate this effect is not very precisely estimated. For imports the corresponding coefficient is larger in absolute value and more accurately esti-

---

[6]As side benefit, our empirical specification of gravity theory delivers estimates of the elasticity of substitution. Due to the use of time dummies this coefficient is commonly not uniquely identified in the standard setup.

[7]Baldwin and Taglioni (2005) argue that due to firm entry and exit into the sector of traded goods the true effect of exchange rate uncertainty on trade is non-linear. Indeed, Herwartz and Weber (2005) find evidence for non-linearities in the effect of exchange rate uncertainty on trade growth.

mated indicating that if domestic goods become more expensive relative to a weighted basket of foreign goods this fosters imports significantly (*reer* is the same regressor regardless of the dependent variable being exports or imports). Real bilateral exchange rates *rex* between reporter and partner country do not add significant information suggesting that real effective exchange rates *reer* already reflect bilateral variation to a sufficient degree.

In contrast, real bilateral exchange rates with the U.S. *rexus* matter with a positive coefficient. This lends support to the expenditure switching hypothesis discussed above: After January 1999 the euro fell sharply so that products sold in euro became cheaper relative to products sold in U.S. dollar. As a consequence part of European demand for foreign products was redirected to European products fostering exports and imports among European countries. Finally, the effect of energy prices is statistically significant over all trade relationships and for each subset individually. High energy prices as a measurable proxy for transportation costs accordingly reduce exports and imports.

6.2 Timing, Speed and Size of Transition

The literature on the trade effect of the euro relies almost exclusively on specifying an instantaneous and exhaustive level shift in January 1999. Accordingly, there is a predominant focus on a single dimension of euro transition, namely size.[8] The smooth transition path estimated here allows to discriminate three dimensions of adjustment, namely timing, speed of adjustment and size. Each dimension is represented by one parameter, $\zeta_k^{(s)}$, $\theta_{k1}^{(s)}$ and $\theta_{k0}^{(s)}$, respectively.

Figure 1 shows the percentage change in trade that is due to the estimated smooth transition path in trade costs. We report results directly in terms of the percentage change in trade rather than trade costs since this corresponds to the convention in the literature. In particular, the figure shows the weighted average transition path for the subsets of trade relationships U2 and O2 (first column) and for the subsets IN and OUT (second column). The top row refers to import data whereas the bottom row refers to exports. Weighted average transition paths are averages over transition paths specific to each sector in each trade relationship based on estimates of $\zeta_k^{(s)}$, $\theta_{k1}^{(s)}$ and $\theta_{k0}^{(s)}$. Corresponding weights are relative sector size $w_k^{(s)}$ normalized to sum to unity for each subset.

---

[8]Some studies make inference on timing of the euro effect by interacting consecutive time dummies with a euro dummy. One potential flaw of such interaction terms is that they confuse omitted trade costs with the euro effect. This appears particularly likely when the specification relies on one set of time dummies common to all trade relationships and thus fails to control for trade relationship specific unobserved trade costs as suggested by theory.

Our main finding is evident from figure 1. Trade within the euro area (U2) increases by 10 to 20 percent between the years 2000 and 2003 compared with trade between European countries that are not members of the euro area (O2). We now discuss each dimension of the transition path.

Probably the most remarkable feature of the figure is that exports start their transition one year *after* the introduction of the single currency in January 1999. Interestingly, this timing pattern also surfaces for exports that enter and leave the euro area as visible from the IN and OUT paths in the second column of the figure. Our results thus contrast the widespread convention to assume a break point in European trade costs in 1999. When comparing our approach to one that relies on a euro dummy one should take into account that most studies on the trade effect of the euro use annual data. In contrast, we employ monthly data to collect as much information around the hypothesized break point as possible. Even with annual data, however, it will be difficult to reconcile the timing pattern identified here with the conventional euro dummy specification.

Evidently, our results are also not consistent with the anticipatory increase in euro area trade before January 1999 which is reported in Micco, Stein, and Ordonez (2003) and Flam and Nordström (2006b). One interpretation of this difference is that anticipation effects in fact are spurious findings due to unaccounted time variation in trade relationship or sector specific trade costs. Moreover, according to our estimates there is not much scope left for long run effects to develop.

Finally, we identify a timing pattern for import data in the first row of figure 1, which is very similar to the pattern identified for export data. The only minor difference is that import data appears less informative. Systematic measurement error in trade figures, which induces a gap between export and import statistics, thus does not seem a relevant alternative explanation for the increase in euro area trade because results for export and import data remain fairly similar both qualitatively and quantitatively. Also, we obtain these estimates while controlling for bilateral and multilateral exchange rate measures to address potential expenditure switching effects due to the euro's depreciation right after its introduction.

Figure 1 also provides insight into the speed of adjustment. It takes about three years for exports and imports to adjust to a new level. This observation applies to trade within the euro area, trade among euro area outsiders and trade between the euro area and euro area outsiders. Compared to adjustment of trade at the aggregate level, adjustment of trade at the sector level is much faster. Aggregate adjustment is more spread out and gradual because different trade sectors adjust at distinct times. Table 2 reports estimates of θ_1 which indicate rapid adjustment within a few month throughout whereas mean estimates of ζ jointly with their standard errors reflect the extent to which differences in

18

timing exist both across trade relationships and across trade sectors. The high speed of adjustment at the sector level fits well with recent microfoundations of the euro's trade effect put forth by Baldwin and Taglioni (2005). These authors argue that trade adjustment to the euro is fast because reduction of exchange rate volatility induces a large number of small firms to enter export markets.

The size of adjustment in figure 1 is the dimension of the transition path that compares most easily with existing studies. At the end of our sample period, mean effects are about 20 (13) percent for export data (imports) when taking the difference between adjustment of U2 versus O2 countries as is common practice. Most of the adjustment is a reduction of trade among O2 countries rather than an increase of trade among U2 countries. In table 3 we obtain essentially identical conclusions about long run effects that prevail once transition is fully completed.

Baldwin (2006a) summarizes the literature on the euro's trade effect as suggesting a boost in trade of about 5 to 10 percent. Thus, our mean estimates are roughly twice as large compared to what has been reported so far even though confidence bands comprise effects of 5 to 10 percent in size. We obtain larger mean estimates for two reasons. First, suppose that mean paths in figure 1 indeed reflect the true transition path. Minimizing the squared error between true transition and a euro dummy delivers smaller estimates because the conventional euro dummy with break point in 1999 kicks in too early. To this end, timing of transition matters for conclusions about the size of transition. Second, table 3 implies a considerably smaller mean difference of long run transition in the homogeneous smooth transition state space model of about 10 (8) percent for exports (imports). Disregarding sectoral heterogeneity thus leads to a downward bias in the extent by which euro area trade is estimated to increases.

Finally, the second column of figure 1 shows adjustment of euro area trade with the United Kingdom, Sweden and Denmark. Exports of these countries into the euro area (IN) fall substantially after 2000 whereas euro area exports to these countries (OUT) slightly dip to reach a higher level thereafter. For import data, IN indicates imports of the United Kingdom, Sweden and Denmark from the euro area (the sets of countries subsumed under IN and OUT remain identical for export and import data). Even though mean estimates decrease at the end of 2001 nothing conclusive follows due to poor estimation precision. In turn, OUT indicates a significant decrease of euro area imports from the United Kingdom, Sweden and Denmark between 2000 and 2003. Overall import and export data suggests that the euro area imports less from but exports more to European countries which have not adopted the euro.

There exists a tight coincidence between U2/O2 transition dynamics and IN/OUT transition dynamics in figure 1. In both cases, transition takes place during the 2000 to 2003

period. However, neither do we impose a restriction on timing coefficients for IN/OUT transition nor do we tie timing coefficients for IN/OUT to those for U2/O2. Therefore, this coincidence in the timing of transition appears an important feature of the data.

Taken together with the observation that the euro area imports less from but exports more to third European countries, one explanation consistent with our estimates is that the euro creates stiffer competition among EMU exporters thereby depressing price markups. Accordingly, part of euro area trade with third countries is redirected back into the euro area. At the same time, third countries import more from the euro area because euro area products become cheaper relative to products in the United Kingdom, Sweden and Denmark.

7 Conclusion

In this paper, we estimate the 'When, how fast and by how much' of adjustment in euro area trade costs during the period from 1995 to 2006. Beyond the flexible smooth transition specification, our approach allows for sector specific impact of trade costs on sectoral trade while controlling for unobserved trade costs and competitiveness terms at the sector level. We find that adjustment in trade costs takes place between the years 2000 and 2003 which suggests that a euro dummy which signifies the introduction of the euro in January 1999 is to some degree misspecified. Adjustment of individual sectors is extremely fast whereas aggregate adjustment is gradual because different sectors adjust at distinct times. These findings support recent microfoundations of the euro's trade effect which predict the effect to happen rapidly.

References

ANDERSON, J. E., AND E. VAN WINCOOP (2003): "Gravity with Gravitas: A Solution to the Border Puzzle," *American Economic Review*, 93(1), 170–192.

——— (2004): "Trade Costs," *Journal of Economic Literature*, 42(3), 691–751.

BAIER, S. L., AND J. H. BERGSTRAND (2001): "The Growth of World Trade: Tariffs, Transport Costs, and Income Similarity," *Journal of International Economics*, 53, 1–27.

BALDWIN, R. E. (2006a): "The Euro's Trade Effects," ECB Working Paper 594.

——— (2006b): *In or Out: Does It Matter? An Evidence-Based Analysis of the Euro's Trade Effects*. Centre for Economic Policy Research.

BALDWIN, R. E., F. SKUDELNY, AND D. TAGLIONI (2005): "Trade Effects of the Euro: Evidence from Sectoral Data," ECB Working Paper 446.

BALDWIN, R. E., AND D. TAGLIONI (2005): "Positive OCA Criteria: Microfoundations for the Rose Effect," unpublished manuscript.

——— (2006): "Gravity for Dummies and Dummies for Gravity Equations," NBER Working Paper 12516.

BAUWENS, L., M. LUBRANO, AND J.-F. RICHARD (2000): *Bayesian Inference in Dynamic Econometric Models*. Oxford University Press: Oxford.

BERGER, H., AND V. NITSCH (2008a): "Gotcha! A Profile of Smuggling in International Trade," unpublished manuscript.

——— (2008b): "Zooming Out: The Trade Effect of the Euro in Historical Perspective," *Journal of International Money and Finance*, (forthcoming).

BRODA, C., AND D. E. WEINSTEIN (2004): "Globalization and the Gains from Variety," NBER Working Paper 10314.

CHANG, Y., J. I. MILLER, AND J. Y. PARK (2008): "Extracting a Common Stochastic Trend: Theories with Some Applications," *Journal of Econometrics*, (forthcoming).

DENARDIS, S., AND C. VICARELLI (2007): "The Impact of the Euro on Trade. The (Early) Impact is not so Large," in *Currencies and Globalization*, ed. by M. L. Ferrington, pp. 123–138. Nova Science Publishers, Inc.

FEENSTRA, R. (2003): *Advanced International Trade: Theory and Evidence*. Princeton University Press: Princeton.

FLAM, H., AND H. NORDSTRÖM (2006a): "Euro Effects on the Intensive and Extensive Margins of Trade," CESIFO Working Paper 1881.

——— (2006b): "Trade Volume Effects of the Euro: Aggregate and Sector Estimates," Seminar Paper 746, Stockholm University.

FULLER, W. (1976): *Introduction to Statistical Time Series*. Wiley: New York.

HERWARTZ, H., AND H. WEBER (2005): "Exchange Rate Uncertainty and Trade Growth – A Comparison of Linear and Nonlinear (Forecasting) Models," *Applied Stochastic Models in Business and Industry*, 21(1), 1–16.

JONES, R. (1985): *Time Series Analysis with Unequally Spaced Data*, vol. 5. Handbook of Statistics.

LUUKKONEN, R., P. SAIKKONEN, AND T. TERÄSVIRTA (1988): "Testing Linearity against Smooth Transition Autoregressive Models," *Biometrika*, 75(3), 491–499.

MICCO, A., E. STEIN, AND G. ORDONEZ (2003): "The Currency Union Effect on Trade: Early Evidence from EMU," *Economic Policy*, 18(37), 315–356.

MONGELLI, F. P., E. DORRUCCI, AND I. AGUR (2005): "What does European Institutional Integration tell us about Trade Integration?," ECB Occasional Paper Series 40.

NITSCH, V., AND M. PISU (2008): "Scalpel, Please! Dissecting the Euros Effect on Trade," unpublished manuscript, Freie Universität Berlin.

PESARAN, M. H., AND R. SMITH (1995): "Estimating Long-Run Relationships from Dynamic Heterogeneous Panels," *Journal of Econometrics*, 68, 79–113.

PHILLIPS, P. C., AND H. R. MOON (1999): "Linear Regression Limit Theory for Nonstationary Panel Data," *Econometrica*, 67(5), 1057–1111.

TAGLIONI, D. (2002): "Exchange Rate Volatility as a Barrier to Trade: New Methodologies and Recent Evidence," *Économie Internationale*, 89(90), 227–259.

WHITE, H. (1980): "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," *Econometrica*, 48(4), 817–838.

A Weighted Moments and Transition

Mean group estimation in dynamic panel models is considered by Pesaran and Smith (1995). In case of panel heterogeneity the mean group estimator measures marginal impacts for an average cross section member. In the spirit of Phillips and Moon (1999) mean group estimation also guards against spurious inferential conclusions that might be attributed to single equation regressions with nonstationary data. As a possible statistical quantity characterizing the parameter heterogeneity one may consider the standard deviation of mean estimates.

Construction of Weights: Weights $w_k^{(s)}$ represent average relative imports (exports) in a given sector k and conditional on trade relationship s where the average is taken over the period 2001:06 to 2006:05,

$$w_k^{(s)} = \frac{\bar{y}_k^{(s)}}{\sum_s \bar{Y}^{(s)}} \quad , \quad \sum_s \sum_k w_k^{(s)} = 1 \,,$$

where $\bar{Y}^{(s)} = \sum_k \bar{y}_k^{(s)}$ with $k = 1,\ldots,K$ and $s = 1,\ldots,S$.

Coefficients and Standard Errors: Let sub denote a subset of the 30 trade relationships we consider, i.e. either ALL, U2, O2, IN or OUT where ALL denotes the full set of trade relationships. Let $\mathcal{S}$ denote the number of trade relationships in sub. Then, the weighted mean group estimator and a corresponding standard error are, respectively,

$$\beta^{(sub)} = \sum_{s \in sub} \sum_k \frac{w_k^{(s)}}{\sum_{s \in sub} \sum_k w_k^{(s)}} \beta_k^{(s)} \text{ and } \omega_\beta^{(sub)} = \left(\frac{1}{\mathcal{S}} \sum_{s \in sub} \sum_k \frac{w_k^{(s)}}{\sum_{s \in sub} \sum_k w_k^{(s)}} (\beta_k^{(s)} - \bar{\beta})^2 \right)^{\frac{1}{2}} .$$

Transition: Smooth transition is a logistic cumulative distribution function (omit s superscript and k subscript), $[1 + \exp\{-\theta_1(t_\kappa/(T\sqrt{0.08\bar{3}}) - \zeta)\}]^{-1} = \kappa$ where κ denotes percent of completed transition at date t_κ. Solve for t_κ,

$$t_\kappa = T\sqrt{0.08\bar{3}} \left(\zeta - \tfrac{1}{\theta_1} \ln(\tfrac{1}{\kappa} - 1) \right) .$$

The number of months needed to complete medium $(1 - 2\alpha)\%$ transition is the difference $t_{1-\alpha} - t_\alpha = 2T\sqrt{0.08\bar{3}}/\theta_1 \ln((1 - \alpha)/\alpha)$. With $\alpha = .05, T = 137$ and $\theta_1 = 232.89$ the fastest 90% of transition happen in one month. The symmetry point of the transition function obtains with $\kappa = 0.5$ as $t_{0.5} = T\sqrt{0.08\bar{3}}\,\zeta$. Converting the lower and upper bound of $\zeta \in [0.30343, 3.186]$ into month delivers [12,126]. To approximate a dummy that kicks in at 1999:01 fix the symmetry point $\zeta = 1.2137$ and set the transition speed to $\theta_1 = 232.89$.

.

B   Kalman Recursions

Given the parameters of the state-space model in (2) to (5), $\psi_k^{(s)\prime}, \phi^{(s)\prime}$, the Kalman filter provides sequentially linear projections for the dynamic system. The likelihood of the model is computed stepwise. In the following reported estimates will have a second index reflecting the time point up to which data for the computations are collected. Such an extra index easily allows to discriminate between forecasts and updates. The analyst is assumed to have some guess concerning the initial states of the system $\left(\text{denoted} \lambda_{k,0|0}^{(s)}\right)$ and their variances $\left(P_{k,0|0}^{(s)}\right)$. The Kalman recursions for regression models with missing observations are given by the following steps (Jones (1985)):

1. Computation of a one step ahead forecast for the state and the associated variance:

$$
\begin{aligned}
\lambda_{k,t|t-1}^{(s)} &= \lambda_{k,t-1|t-1}^{(s)} \\
P_{k,t|t-1}^{(s)} &= P_{k,t-1|t-1}^{(s)} + h_k^{(s)}.
\end{aligned}
$$

2. The forecast of the state and observable explanatory variables are used to obtain a prediction for the dependent variable:

$$
y_{k,t|t-1}^{(s)} = q_{it}^{(s)} \beta_{ik}^{(s)} + q_{jt}^{(s)} \beta_{jk}^{(s)} + (1 - \sigma^{(s)}) \left(\ln(\tau_{k,t}^{(s)}) + \lambda_{k,t|t-1}^{(s)} \right). \tag{8}
$$

3. Comparing $y_{kt}^{(s)}$ and $y_{k,t|t-1}^{(s)}$ is feasible in case that $y_{kt}^{(s)}$ is observed. Then, the prediction error $u_{kt}^{(s)}$ with variance $W_{kt}^{(s)}$ is obtained as:

$$
\begin{aligned}
w_{kt}^{(s)} &= y_{kt}^{(s)} - y_{k,t|t-1}^{(s)} \\
W_{kt}^{(s)} &= (1 - \sigma^{(s)})^2 P_{k,t|t-1}^{(s)} + g_k^{(s)}.
\end{aligned}
$$

4. The latter quantities contribute to the models' log likelihood with

$$
l_{kt}^{(s)} = -0.5 \ln(2\pi) - 0.5 \left(u_{kt}^{(s)} \right)^2 / W_{kt}^{(s)} - 0.5 \ln W_{kt}^{(s)}. \tag{9}
$$

5. The innovation u_{kt} and its variance are used to update the current estimate of the state vector:

$$
\begin{aligned}
\lambda_{k,t|t}^{(s)} &= \lambda_{k,t|t-1}^{(s)} + P_{k,t|t-1}^{(s)} (1 - \sigma^{(s)}) u_{kt}^{(s)} / W_{kt}^{(s)} \\
P_{k,t|t}^{(s)} &= P_{k,t|t-1}^{(s)} + (1 - \sigma^{(s)})^2 \left(P_{k,t|t-1}^{(s)} \right)^2 / W_{kt}^{(s)}.
\end{aligned}
$$

Note that the log likelihood function integrates over all time and sector specific estimates l_{kt} given in (9), i.e.

$$l = l\left(\boldsymbol{\psi}_k^{s\prime}, \boldsymbol{\phi}^{s\prime}\right)' = \sum_k \sum_t l_{kt}.$$

In case a particular observation on $y_{kt}^{(s)}$ is missing, steps 3. and 4. are left out and the updating in step 5. becomes

$$
\begin{aligned}
\lambda_{k,t|t}^{(s)} &= \lambda_{k,t|t-1}^{(s)} \\
P_{k,t|t}^{(s)} &= P_{k,t|t-1}^{(s)}.
\end{aligned}
$$

C Serial Correlation Tests

Serial correlation might easily be diagnosed by means of Portmanteau type test statistics exploiting the autocorrelation coefficients of the estimated model residuals $\hat{u}_{kt}^{(s)}$. To obtain an indication of serially correlated error terms which is robust under heteroskedasticity, however, we rather use the following auxiliary regression:

$$\hat{u}_{kt}^{(s)} = c + \kappa_1 \hat{u}_{kt-1}^{(s)} + \ldots + \kappa_h \hat{u}_{kt-h}^{(s)} + v_t,$$

where c is an intercept term and v_t a white noise disturbance. We test the null hypothesis $H_0 : \kappa_1 = \kappa_2 = \ldots = \kappa_h = 0$ by means of a Wald-test

$$\omega_h = \hat{\kappa}'(\text{Cov}[\hat{\kappa}])^{-1}\hat{\kappa} \xrightarrow{d} \chi^2(h). \tag{10}$$

To implement the statistic in (10) we use the heteroskedasticity consistent covariance estimator for the estimated parameter vector $\hat{\kappa} = (\hat{\kappa}_1, \hat{\kappa}_2, \ldots, \hat{\kappa}_h)'$ (White (1980)). With respect to the choice of the lag order h we consider tests on serial correlation at lag 1 and joint correlation at lags 1 to 12. The latter choices appear reasonable noting that monthly data enter our analysis.

D Data Appendix

We seasonally adjust trade data by means of seasonal dummies. Merging data in value and in volume allows to express exports and imports in constant prices of 2000. To do so, we compute implicit unit price deflators and use the average of the 12 price observations in 2000 to re-value volumes. Monthly industrial production data comes from International Financial Statistics (IFS) of the IMF. Monthly exchange rates are market rates from

IFS. Daily exchange rate data used to compute exchange rate volatility comes from the FED historical database.[9] The IFS indicators of real effective exchange rates based on unit labor costs in manufacturing represent the product of the index of the ratio of the relevant indicator (in national currency) for the country listed to a weighted geometric average of the corresponding indicators for 20 other industrial countries. Bilateral real exchange rates are computed as $rex = eP_{par}/P_{rep}$ where e is reporter's currency in terms of partner's currency and P_{rep} (P_{par}) denotes reporter's (partner's) producer price index. For $rexus$ the partner country is the U.S.. Producer price indices and the energy price index are drawn from IFS. All indices are normalized to a base year 2000. We take natural logs of all series unless otherwise noted.

---

[9]http://www.federalreserve.gov/releases/h10/Hist/default1999.htm
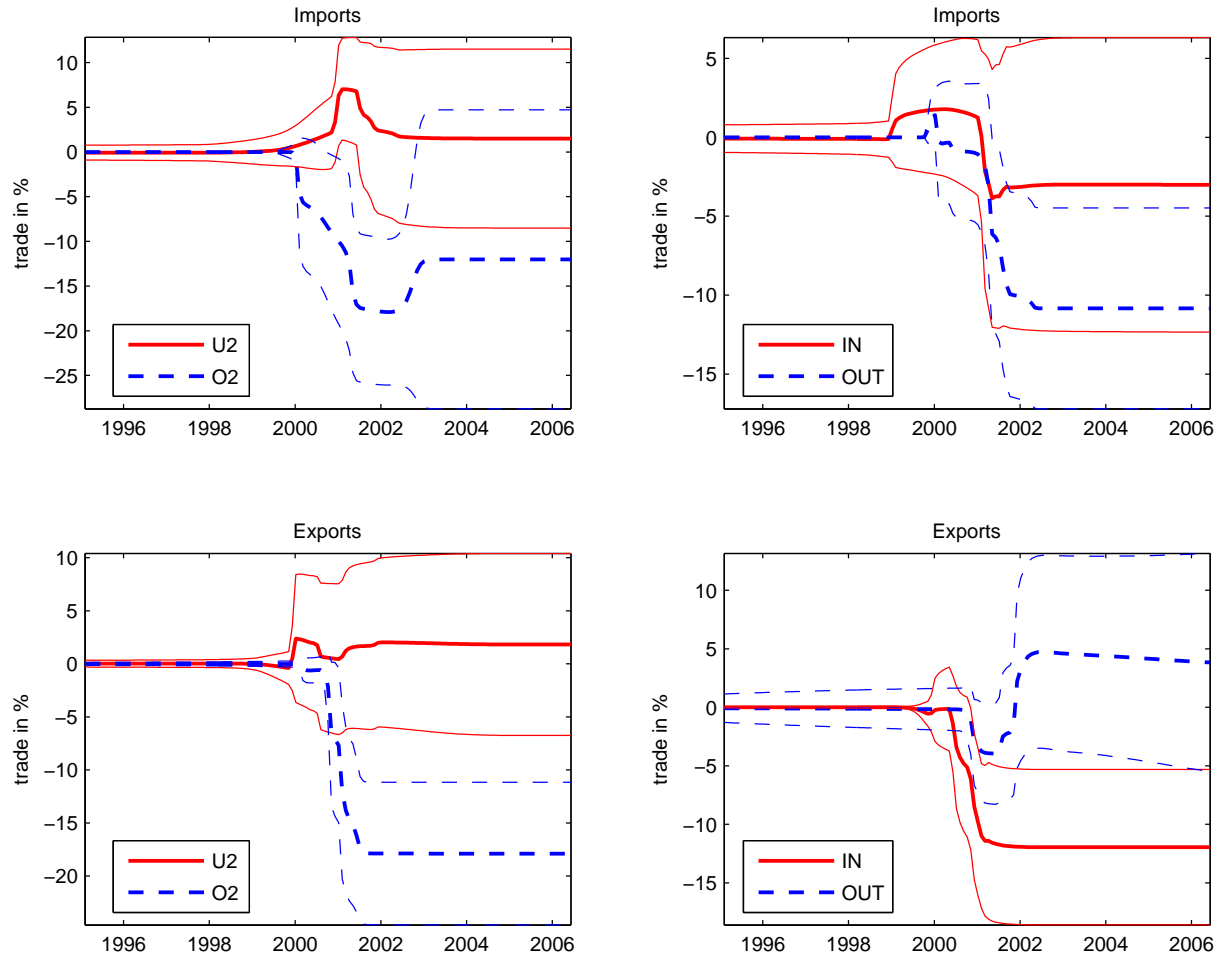
Figure 1: Weighted averages of sector specific transition paths of U2 and O2 countries (first column) and of IN and OUT countries (second column). All estimates are based on the functional smooth transition state space model. The first row refers to import data. The second row refers to export data. Confidence bands are based on a 10% significance level (see appendix A for computation).

Table 1: LR Specification Tests and Model Diagnostics

i,j		Imports						Exports					
		LR_d	SR	LR_f	AR1	$I(1)$	LR_X	LR_d	SR	LR_f	AR1	$I(1)$	LR_X
GER,FRA	U2	5.11	5.1	7999.2	.33	.98	6.56	2.94	5.2	6240.5	.35	.98	17.7
GER,ITA	U2	6.45	4.5	11153.	.30	.99	12.8	3.74	4.8	6269.5	.31	.97	5.85
GER,UK	OUT	5.49	3.8	2868.1	.28	.96	7.31	4.35	4.1	4884.7	.31	.99	10.6
GER,SWE	OUT	4.14	3.2	2965.5	.17	.90	2.66	5.05	3.2	2260.9	.32	.93	9.04
GER,DK	OUT	2.32	3.4	4958.5	.31	.93	21.5	6.71	3.6	1668.5	.20	.98	10.5
FRA,GER	U2	8.44	7.5	5916.7	.25	.91	29.4	5.71	7.5	6939.3	.57	.90	33.6
FRA,ITA	U2	4.91	6.2	5279.1	.67	.83	47.6	2.87	6.0	4635.4	.39	.94	29.2
FRA,UK	OUT	2.67	5.3	5861.5	.25	.95	34.7	1.83	5.2	7747.5	.37	.95	52.7
FRA,SWE	OUT	1.60	3.5	2051.4	.25	.93	14.5	5.72	3.7	2553.2	.40	.90	16.8
FRA,DK	OUT	1.85	3.3	1991.8	.27	.91	25.7	3.61	3.6	3619.9	.34	.87	21.6
ITA,GER	U2	4.45	6.2	3028.8	.22	.92	30.2	4.95	6.7	6418.9	.46	.87	18.9
ITA,FRA	U2	3.79	5.6	3662.9	.41	.85	10.4	10.1	5.9	5658.9	.74	.80	20.6
ITA,UK	OUT	4.85	4.1	4023.7	.29	.90	50.7	2.63	4.2	4769.6	.37	.83	15.7
ITA,SWE	OUT	4.04	3.3	1790.1	.29	.92	14.3	1.70	3.3	1842.1	.53	.80	53.6
ITA,DK	OUT	10.3	3.1	734.6	.31	.92	16.7	3.25	3.4	2251.0	.41	.84	18.4
UK,GER	IN	5.05	4.0	2883.9	.21	.96	9.44	5.79	4.6	5122.7	.22	.97	25.9
UK,FRA	IN	3.47	4.0	3293.7	.20	.95	17.4	1.85	4.9	3382.7	.19	.98	8.53
UK,ITA	IN	3.69	3.5	3721.6	.19	.95	18.2	6.68	3.8	3337.3	.14	.94	10.6
UK,SWE	O2	0.83	3.3	1627.6	.22	.97	62.8	4.75	3.4	3029.2	.27	.93	3.02
UK,DK	O2	1.34	2.7	963.5	.27	.83	13.0	5.27	2.9	1280.7	.27	.90	7.98
SWE,GER	IN	1.45	4.5	2814.4	.30	.97	15.0	0.02	3.8	3125.5	.38	.91	19.6
SWE,FRA	IN	0.98	3.3	2145.4	.32	.99	11.1	0.90	3.6	1428.6	.35	.91	6.81
SWE,ITA	IN	1.84	3.1	2316.9	.44	.84	5.83	2.10	3.5	1585.3	.26	1.0	8.05
SWE,UK	O2	1.23	3.3	2262.3	.29	.94	50.8	0.11	3.4	1122.0	.40	.94	9.62
SWE,DK	O2	6.41	4.7	5520.2	.15	.98	46.5	2.57	4.1	3129.7	.26	.95	11.8
DK,GER	IN	4.13	4.0	4219.6	.93	.98	46.2	1.41	4.9	3630.2	.61	.94	36.9
DK,FRA	IN	3.39	3.1	2409.9	.33	.84	11.8	6.83	3.3	997.8	.21	.98	14.4
DK,ITA	IN	1.07	3.4	3193.7	.24	.97	3.60	1.24	3.2	635.7	.32	1.0	15.4
DK,UK	O2	0.56	3.1	2319.5	.22	.94	16.1	4.21	3.3	797.7	.30	.97	13.4
DK,SWE	O2	14.7	4.0	3194.4	.28	.97	10.9	19.7	4.9	5587.4	.20	1.0	5.94
crit	5%	5.99	-	14.07	.05	-	11.1	5.99	-	14.07	.05	-	11.1
	10%	4.60	-	12.02		-	9.24	4.60	-	12.02		-	9.24

Notes: LR_d and LR_f are LR statistics testing the smooth transition homogeneous state space model ($\boldsymbol{\psi}_1^{(s)} = 0$) against a euro dummy homogeneous state space model and the functional smooth transition state space model, respectively. Conditional on the functional state space model, LR_X measures the explanatory content of five additional exogenous variables. SR denotes the ratio of standard error estimates obtained when excluding unobserved variables $\lambda_{kt}^{(s)} = 0$ over standard error estimates from the homogeneous state space model. Estimated residuals of the observation equation (2) of the functional state space model excluding exogenous control variables ($\boldsymbol{\psi}_1^{(s)} \neq 0$, $\boldsymbol{\gamma} = 0$) are diagnosed for stationarity ($I(1)$) and first order serial correlation (AR1). Diagnostics AR1 and $I(1)$ are frequencies of rejections of the null hypothesis over sectors. 'U2', 'O2', 'OUT' and 'IN' classify trade relationships.

Table 2: Mean Group Estimates

Imports	ALL coef.	ALL std.	U2 coef.	U2 std.	O2 coef.	O2 std.	IN coef.	IN std.	OUT coef.	OUT std.
σ	5.016	0.072	4.949	0.121	5.458	0.082	5.074	0.099	5.176	0.067
c	7.161	1.260	8.386	2.583	8.412	2.323	4.497	2.195	7.236	2.937
β_i	0.957	0.169	1.067	0.267	0.901	0.510	1.184	0.351	0.419	0.359
β_j	0.215	0.130	0.080	0.319	0.205	0.448	0.336	0.173	0.408	0.204
g	0.499	0.042	0.376	0.074	0.736	0.067	0.613	0.047	0.624	0.061
h	0.048	0.007	0.036	0.004	0.058	0.006	0.046	0.002	0.043	0.003
θ_0	0.042	0.037	0.000	0.017	0.038	0.023	0.010	0.015	0.031	0.010
θ_1	203.3	12.91	199.2	30.37	182.6	33.23	220.0	15.46	231.4	4.056
ζ	1.854	0.053	1.830	0.080	1.836	0.099	1.820	0.070	1.769	0.078
vol	0.038	0.021	0.084	0.051	0.084	0.030	-0.056	0.023	0.053	0.025
$reer$	0.488	0.165	0.347	0.369	-0.143	0.279	0.962	0.234	0.344	0.349
rex	0.006	0.161	-0.373	0.438	0.012	0.189	0.453	0.125	0.261	0.230
$rexus$	0.249	0.068	0.287	0.156	0.297	0.277	0.164	0.095	0.306	0.107
en	-0.051	0.014	-0.062	0.025	-0.052	0.021	-0.009	0.024	-0.068	0.030
Exports										
σ	5.345	0.077	5.494	0.168	5.402	0.230	5.314	0.095	5.090	0.113
c	8.812	1.195	11.571	2.995	10.617	1.749	5.528	1.304	5.982	1.531
β_i	0.871	0.126	0.914	0.244	0.306	0.308	1.043	0.239	0.821	0.237
β_j	0.637	0.157	0.490	0.248	-0.123	0.311	0.824	0.323	0.968	0.322
g	0.469	0.041	0.334	0.079	0.712	0.062	0.551	0.063	0.576	0.055
h	0.042	0.005	0.039	0.016	0.060	0.006	0.046	0.004	0.039	0.002
θ_0	0.008	0.009	-0.001	0.013	0.046	0.012	0.031	0.010	-0.003	0.024
θ_1	204.4	11.56	188.96	30.00	196.7	24.31	214.6	19.09	226.0	11.85
ζ	1.772	0.047	1.682	0.109	1.776	0.047	1.718	0.044	1.973	0.078
vol	0.014	0.015	0.024	0.031	0.029	0.022	-0.039	0.030	0.033	0.024
$reer$	-0.208	0.128	-0.405	0.306	0.129	0.220	0.088	0.148	-0.186	0.233
rex	0.219	0.155	0.227	0.333	0.022	0.192	0.520	0.408	0.025	0.164
$rexus$	0.131	0.054	0.050	0.135	0.251	0.103	0.208	0.052	0.180	0.094
en	-0.030	0.012	-0.032	0.023	-0.031	0.029	-0.041	0.020	-0.017	0.027

Notes: Estimates are based on the functional smooth transition state space model. See appendix A for computation of weighted average coefficients 'coef.' and corresponding standard errors 'std.'. For the set of exogenous variables $vol, reer, rex, rexus$ and en the table records the weighted average of the element in $\boldsymbol{\gamma}_k^{(s)}$, which relates to the respective variable, multiplied by $(1 - \sigma^{(s)})$. These estimates thus reflect the elasticity of trade with respect to the particular exogenous variable.

Table 3: Long Run Adjustment as Change of Imports and Exports in Percent

	U2		O2		IN		OUT	
Imports	coef.	std.	coef.	std.	coef.	std.	coef.	std.
Func. Model	1.435	6.131	-12.01	10.14	-3.064	5.708	-10.84	3.857
Hom. Model	-2.958	5.786	-11.04	7.561	6.021	4.779	-7.684	4.713
Exports								
Func. Model	1.847	5.213	-17.90	4.077	-11.96	4.028	3.547	5.998
Hom. Model	0.686	4.683	-8.988	7.505	-13.66	2.945	-4.443	4.121

Notes: Reported statistics are weighted averages of $\exp\{(1 - \sigma^{(s)})\theta_{k0}^{(s)}\} - 1$, i.e. weighted average percentage change in exports (imports) after fully completed transition. We then use weights $w_k^{(s)}$ accordingly normalized to compute averages in the table. Weighted averages are based on the indicated subset of $\theta_{k0}^{(s)}$ coefficients. 'coef.' abbreviates coefficient estimate and 'std' denotes the standard error. For computation of these statistic see appendix A.

SFB 649 Discussion Paper Series 2008

001 "Testing Monotonicity of Pricing Kernels" by Yuri Golubev, Wolfgang Härdle and Roman Timonfeev, January 2008.

002 "Adaptive pointwise estimation in time-inhomogeneous time-series models" by Pavel Cizek, Wolfgang Härdle and Vladimir Spokoiny, January 2008.

003 "The Bayesian Additive Classification Tree Applied to Credit Risk Modelling" by Junni L. Zhang and Wolfgang Härdle, January 2008.

004 "Independent Component Analysis Via Copula Techniques" by Ray-Bing Chen, Meihui Guo, Wolfgang Härdle and Shih-Feng   Huang, January 2008.

005 "The Default Risk of Firms Examined with Smooth Support Vector Machines" by Wolfgang Härdle, Yuh-Jye Lee, Dorothea Schäfer and Yi-Ren Yeh, January 2008.

006 "Value-at-Risk and Expected Shortfall when there is long range dependence" by Wolfgang Härdle and Julius Mungo, Januray 2008.

007 "A Consistent Nonparametric Test for Causality in Quantile" by Kiho Jeong and Wolfgang Härdle, January 2008.

008 "Do Legal Standards Affect Ethical Concerns of Consumers?" by Dirk Engelmann and Dorothea Kübler, January 2008.

009 "Recursive Portfolio Selection with Decision Trees" by Anton Andriyashin, Wolfgang Härdle and Roman Timofeev, January 2008.

010 "Do Public Banks have a Competitive Advantage?" by Astrid Matthey, January 2008.

011 "Don't aim too high: the potential costs of high aspirations" by Astrid Matthey and Nadja Dwenger, January 2008.

012 "Visualizing exploratory factor analysis models" by Sigbert Klinke and Cornelia Wagner, January 2008.

013 "House Prices and Replacement Cost: A Micro-Level Analysis" by Rainer Schulz and Axel Werwatz, January 2008.

014 "Support Vector Regression Based GARCH Model with Application to Forecasting Volatility of Financial Returns" by Shiyi Chen, Kiho Jeong and Wolfgang Härdle, January 2008.

015 "Structural Constant Conditional Correlation" by Enzo Weber, January 2008.

016 "Estimating Investment Equations in Imperfect Capital Markets" by Silke Hüttel, Oliver Mußhoff, Martin Odening and Nataliya Zinych, January 2008.

017 "Adaptive Forecasting of the EURIBOR Swap Term Structure" by Oliver Blaskowitz and Helmut Herwatz, January 2008.

018 "Solving, Estimating and Selecting Nonlinear Dynamic Models without the Curse of Dimensionality" by Viktor Winschel and Markus Krätzig, February 2008.

019 "The Accuracy of Long-term Real Estate Valuations" by Rainer Schulz, Markus Staiber, Martin Wersing and Axel Werwatz, February 2008.

020 "The Impact of International Outsourcing on Labour Market Dynamics in Germany" by Ronald Bachmann and Sebastian Braun, February 2008.

021 "Preferences for Collective versus Individualised Wage Setting" by Tito Boeri and Michael C. Burda, February 2008.

022 "Lumpy Labor Adjustment as a Propagation Mechanism of Business Cycles" by Fang Yao, February 2008.

023 "Family Management, Family Ownership and Downsizing: Evidence from S&P 500 Firms" by Jörn Hendrich Block, February 2008.

024 "Skill Specific Unemployment with Imperfect Substitution of Skills" by Runli Xie, March 2008.

025 "Price Adjustment to News with Uncertain Precision" by Nikolaus Hautsch, Dieter Hess and Christoph Müller, March 2008.

026 "Information and Beliefs in a Repeated Normal-form Game" by Dietmar Fehr, Dorothea Kübler and David Danz, March 2008.

027 "The Stochastic Fluctuation of the Quantile Regression Curve" by Wolfgang Härdle and Song Song, March 2008.

028 "Are stewardship and valuation usefulness compatible or alternative objectives of financial accounting?" by Joachim Gassen, March 2008.

029 "Genetic Codes of Mergers, Post Merger Technology Evolution and Why Mergers Fail" by Alexander Cuntz, April 2008.

030 "Using R, LaTeX and Wiki for an Arabic e-learning platform" by Taleb Ahmad, Wolfgang Härdle, Sigbert Klinke and Shafeeqah Al Awadhi, April 2008.

031 "Beyond the business cycle – factors driving aggregate mortality rates" by Katja Hanewald, April 2008.

032 "Against All Odds? National Sentiment and Wagering on European Football" by Sebastian Braun and Michael Kvasnicka, April 2008.

033 "Are CEOs in Family Firms Paid Like Bureaucrats? Evidence from Bayesian and Frequentist Analyses" by Jörn Hendrich Block, April 2008.

034 "JBendge: An Object-Oriented System for Solving, Estimating and Selecting Nonlinear Dynamic Models" by Viktor Winschel and Markus Krätzig, April 2008.

035 "Stock Picking via Nonsymmetrically Pruned Binary Decision Trees" by Anton Andriyashin, May 2008.

036 "Expected Inflation, Expected Stock Returns, and Money Illusion: What can we learn from Survey Expectations?" by Maik Schmeling and Andreas Schrimpf, May 2008.

037 "The Impact of Individual Investment Behavior for Retirement Welfare: Evidence from the United States and Germany" by Thomas Post, Helmut Gründl, Joan T. Schmit and Anja Zimmer, May 2008.

038 "Dynamic Semiparametric Factor Models in Risk Neutral Density Estimation" by Enzo Giacomini, Wolfgang Härdle and Volker Krätschmer, May 2008.

039 "Can Education Save Europe From High Unemployment?" by Nicole Walter and Runli Xie, June 2008.

040 "Solow Residuals without Capital Stocks" by Michael C. Burda and Battista Severgnini, August 2008.

041 "Unionization, Stochastic Dominance, and Compression of the Wage Distribution: Evidence from Germany" by Michael C. Burda, Bernd Fitzenberger, Alexander Lembcke and Thorsten Vogel, March 2008

042 "Gruppenvergleiche bei hypothetischen Konstrukten – Die Prüfung der Übereinstimmung von Messmodellen mit der Strukturgleichungs-methodik" by Dirk Temme and Lutz Hildebrandt, June 2008.

043 "Modeling Dependencies in Finance using Copulae" by Wolfgang Härdle, Ostap Okhrin and Yarema Okhrin, June 2008.

044 "Numerics of Implied Binomial Trees" by Wolfgang Härdle and Alena Mysickova, June 2008.
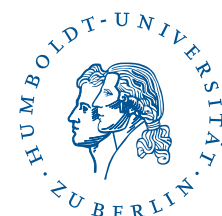
045 "Measuring and Modeling Risk Using High-Frequency Data" by Wolfgang Härdle, Nikolaus Hautsch and Uta Pigorsch, June 2008.

046 "Links between sustainability-related innovation and sustainability management" by Marcus Wagner, June 2008.

047 "Modelling High-Frequency Volatility and Liquidity Using Multiplicative Error Models" by Nikolaus Hautsch and Vahidin Jeleskovic, July 2008.

048 "Macro Wine in Financial Skins: The Oil-FX Interdependence" by Enzo Weber, July 2008.

049 "Simultaneous Stochastic Volatility Transmission Across American Equity Markets" by Enzo Weber, July 2008.

050 "A semiparametric factor model for electricity forward curve dynamics" by Szymon Borak and Rafał Weron, July 2008.

051 "Recurrent Support Vector Regreson for a Nonlinear ARMA Model with Applications to Forecasting Financial Returns" by Shiyi Chen, Kiho Jeong and Wolfgang K. Härdle, July 2008.

052 "Bayesian Demographic Modeling and Forecasting: An Application to U.S. Mortality" by Wolfgang Reichmuth and Samad Sarferaz, July 2008.

053 "Yield Curve Factors, Term Structure Volatility, and Bond Risk Premia" by Nikolaus Hautsch and Yangguoyi Ou, July 2008.

054 "The Natural Rate Hypothesis and Real Determinacy" by Alexander Meyer-Gohde, July 2008.

055 "Technology sourcing by large incumbents through acquisition of small firms" by Marcus Wagner, July 2008.

056 "Lumpy Labor Adjustment as a Propagation Mechanism of Business Cycle" by Fang Yao, August 2008.

057 "Measuring changes in preferences and perception due to the entry of a new brand with choice data" by Lutz Hildebrandt and Lea Kalweit, August 2008.

058 "Statistics E-learning Platforms: Evaluation Case Studies" by Taleb Ahmad and Wolfgang Härdle, August 2008.

059 "The Influence of the Business Cycle on Mortality" by Wolfgang H. Reichmuth and Samad Sarferaz, September 2008.

060 "Matching Theory and Data: Bayesian Vector Autoregression and Dynamic Stochastic General Equilibrium Models" by Alexander Kriwoluzky, September 2008.

061 "Eine Analyse der Dimensionen des Fortune-Reputationsindex" by Lutz Hildebrandt, Henning Kreis and Joachim Schwalbach, September 2008.

062 "Nonlinear Modeling of Target Leverage with Latent Determinant Variables – New Evidence on the Trade-off Theory" by Ralf Sabiwalsky, September 2008.

063 "Discrete-Time Stochastic Volatility Models and MCMC-Based Statistical Inference" by Nikolaus Hautsch and Yangguoyi Ou, September 2008.

064 "A note on the model selection risk for ANOVA based adaptive forecasting of the EURIBOR swap term structure" by Oliver Blaskowitz and Helmut Herwartz, October 2008.

065 "When, How Fast and by How Much do Trade Costs change in the EURO Area?" by Helmut Herwartz and Henning Weber, October 2008.