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Rose Effect and the Euro: The Magic is Gone

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Abstract:
This paper presents an updated meta-analysis of the effect of currency unions on trade, focusing on the Euro area. Using meta-regression methods such as funnel asymmetry test, evidence for strong publication bias is found. The estimated underlying effect for non-Euro studies reaches about 50%. However, the Euro's trade promoting effect corrected for publication bias is insignificant. The Rose effect literature shows signs of the economics research cycle: reported t-statistic is a quadratic function of publication year. Explanatory meta-regression (robust fixed effects and random effects) suggests that some authors produce predictable results. Interestingly, proxies for authors' IT skills were also found significant.

Keywords: Rose effect; Trade; Currency union; Euro; Meta-analysis; Publication bias

JEL: C42; F15; F33

1 Introduction

"Most of the Rose effect literature treats currency unions as magic wands—one touch and intra-currency-union trade flows rise between 5% and 1400%. The only question is: How big is the magic?" (Baldwin 2006, p. 36)

Since the pioneering work of Rose (2000) and his legendary result that currency unions increase trade by more than 200%, a whole new stream of literature has emerged and thrived, in recent years focusing especially on the Eurozone as the most ambitious project of monetary union. How much does the Euro boost trade among the Eurozone members? While some researchers are rather skeptical to search for "the one number" (e.g., Richard Baldwin, as the opening quotation suggests), the others keep seeking: in a narrative literature review, Frankel (2008b) estimates the Euro's Rose effect to lie between 10% and 15%. Even Baldwin (2006, p. 48) himself talks about 5%–10% and expects the effect to double as the Euro matures. This question is very attractive for welfare economists and policy makers: for instance, Frankel (2008a) uses his estimates to give Central- and Eastern-European countries advice on the timing of their admission to the Eurozone; and Masson (2008), employing the result that "currency unions double trade," asseses the welfare effects of creating a monetary union in Africa.

The purpose of this work is to extend the meta-analysis (for an excellent introduction to this methodology and application in economics, see Stanley 2001) of Rose & Stanley (2005) by new studies and new methods, which enables us to concentrate on the effects of the Euro and other currency unions separately. It is shown that this distinction is important since both sub-samples tell a completely different story. Twenty-seven new studies were added to the sample, 21 of which focusing on the Eurozone. Together, there are 61 studies, 28 on the Eurozone and 33 on other currency unions. First, simple fixed and random effects meta-analysis is performed. Publication bias (Card & Krueger 1995; Stanley 2005a) is accounted for by the straightforward "trim and fill" method (Duval & Tweedie 2000); subsequently, impact and citations weights are used and the respective results are compared to the benchmark case. We also check the robustness of the estimates with respect to the most influential studies. The following part examines publication bias among the literature, using the meta-regression approach (Stanley & Jarrell 1989; Stanley *et al.* 2008) and graphical methods (funnel plots, Galbraith plots); the "true" effect is estimated as well. Multiple meta-regression methods, including robust meta-regression (Havránek & Iršová 2008) and random effects meta-regression, are used to model heterogeneity present in the sample. A test for the "economics research cycle" is conducted (novelty and fashion in economics research, see Goldfarb 1995).

The paper is structured as follows: in Section 2, simple fixed and random effects meta-analysis is performed and the basic properties of the sample of literature are described. Section 3 focuses on publication selection and search for the true Rose effect. In Section 4, explanatory meta-regression is conducted. Section 5 concludes.

2 Simple Meta-Analysis

The "Rosean" stream of literature usually employs a variation of the following regression to estimate the trade effect of currency unions, the so-called gravity equation (for a detailed discussion and criticism, see Baldwin 2006):

$$\log T_{ijt} = \alpha_0 + \gamma CU_{ijt} + \chi_1(\log Y_i \log Y_j)_t + \chi_2 \log D_{ij} + \sum_{k=1}^{K} \eta_k X_{ijt} + \epsilon_{ijt}, \quad (1)$$

where T_{ijt} stands for the trade flow between two countries (i and j) in period t, CU is a dummy which equals one if both countries are engaged in a currency union in period t, Y denotes the real GDP, D is the distance between the two countries, and X denotes other control variables. The actual boost to trade due to the formation of a monetary union is thus given by $\gimel \doteq e^{\gamma} - 1$.

The meta-analytical process starts with a selection of literature to be included in the analysis. It is usually advised to use only one estimate from each study since otherwise a single researcher could easily dominate the survey (Stanley 2001; Krueger 2003). This, of course, raises the problem of selecting the "representative" estimate. The present paper builds on the dataset provided by Rose & Stanley (2005) which covers a sample of results taken from 34 papers on currency unions' trade effect. The dataset, however, contains only 7 studies on the Eurozone, which does not make it possible to estimate the Euro's effect separately. For this reason, additional search was conducted in the RePEc and Google Scholar databases, concentrating mainly on new studies estimating the effect of the Euro.[1] All papers on the Rose effect containing a quantitative estimate of γ were included, both published and unpublished, extending the sample to the total of 61 studies, including 28 studies on the Eurozone. The authors' preferred estimates were selected; in case there was no preference expressed, the model with the best fit was chosen. However, most authors in this sample reveal their preferences concerning the "best" estimate directly in the abstract or conclusion.

Meta-analysis has its roots in psychology and epidemiology after the second world war (for an extensive introduction, see Borenstein *et al.* 2009). Originally, it was used to increase the number of observations and thus statistical power in those fields of medical research where experiments were extremely costly and scarce, or to estimate the "true" effect when the findings were seemingly mixed. Subsequently, this method spread to other sciences, including economics (beginning with Stanley & Jarrell 1989). The essence of meta-analysis is to use all available studies since even biased and misspecified results may carry useful information which can be decoded by the meta-regression approach. Omitting some papers *ex ante*, as Baldwin (2006) suggests in his narrative review, is thus the exact opposite of what meta-analysts should do. *"He (Richard Baldwin) thinks he knows which of the studies are good and which are bad (...), and wants only to count the good ones. The problem with this is that other authors have other opinions as to what is good and what is bad."* (Frankel 2006, p. 83). For this reason, we let the data speak for themselves. Fortunately, the meta-

[1] The exact search query used in RePEc was (((currency | monetary) + union) | euro) + trade + (effect | rose) + estimate, abstract search since 2002. The "old" (Rose & Stanley 2005) data were updated—for example, many of the then working papers have been published in a journal since 2005 and their estimates might have slightly changed.

regression methods are able to cope with some degree of misspecification bias (Stanley 2008).

It is generally recognized that the reported Rose effect of the Euro is significantly lower than that of other currency unions taken as a whole (Micco *et al.* 2003; Frankel 2008b). Frankel (2008b) tests three possible explanations (Euro's youth, bigger size of the Eurozone economies compared to average members of other monetary unions, and reverse causality for the earlier studies), but rejects them one by one. For policy recommendations concerning the Euro, in any case, only the estimates derived from the Eurozone studies should be taken into account. The results of the non-Euro papers, however, are useful as well: on the one hand, these studies can serve as a control group; on the other hand, the "true" general Rose effect of other currency unions can be extracted from them.

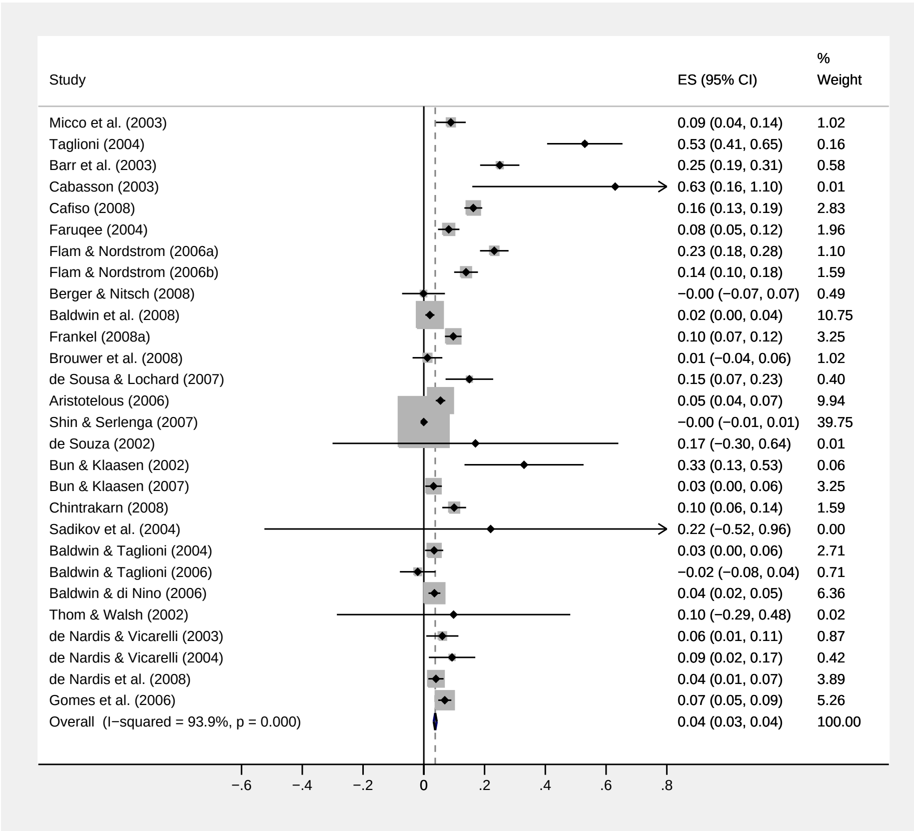

Figure 1: Forest plot of individual estimates, Eurozone studies

The Eurozone sample is depicted in Figure 1; this type of figure is usually called "forest plot" in medical research. Black dots symbolize individual estimates, horizontal lines show the respective 95% confidence intervals. The usual method of combining estimates taken from various studies is the standard fixed effects (FE) estimator[2] which weighs each observation according to its preci-

[2]Note that "fixed" and "random" effects estimators in meta-analysis do not correspond to the standard use of these terms in panel data econometrics. For a more detailed explanation, see Abreu *et al.* (2005) and Sutton *et al.* (2000).

sion; i.e., inverse standard error. The weights constructed on the basis of the inverse-variance method are symbolized by squares with gray fill in the forest plot. The pooled effect estimated by FE is plotted as a vertical dashed line, the solid vertical line symbolizes no effect. Using FE, the pooled estimate of the Euro's γ is very low: a mere 0.038 ($\gimel$ = 3.87%) with 95% confidence interval CI = (3.36%, 4.39%), although it is very significant (z-stat. = 14.9). One caveat with the fixed effects estimator is that it assumes the true effect to be equal accros all studies. However, it is apparent that the confidence intervals in the forest plot often do not overlap; and also the Q-test of heterogeneity rejects the null hypothesis of homogeneity [$\chi^2_{(27)} = 441$, $p < 0.001$]. Therefore, it is more appropriate to use the random effects (RE) estimator which allows the true effect to vary accros studies (thus, rather than the one true effect, RE estimate the average of individual true effects; see Higgins *et al.* 2009, for details). It is also more robust since the weights it assigns to individual studies are more evenly distributed: FE give 39.75% weight to the most influential study (Shin & Serlenga 2007) compared to 4.94% in the case of RE. The respective pooled estimate is 0.092 ($\gimel$ = 9.64%), CI = (6.93%, 12.41%); its significance is lower, but still very high (z-stat. = 7.2).

These results are not very useful for policy purposes, though, because (among other things) they do not account for likely publication selection (preference of editors, referees, or researchers themselves for significant or non-negative results; more on this topic in Section 3). In this univariate framework, there is a remedy proposed by Duval & Tweedie (2000): the nonparametric "trim and fill" method. In a nutshell, the trim and fill method tries to estimate results of studies that are missing in the sample due to publication selection—in this case, 12 more studies are added. Subsequently, fixed and random effects estimators are applied on this unbiased sample. The results are summarized in Table 1: the FE estimator yields 0.26 ($\gimel$ = 2.63%) with a narrow CI, the RE estimator reports 0.035 ($\gimel$ = 3.56%), CI = (0.8%, 6.4%). The Euro's effect on trade is thus still significant, even though the z-statistic decreases to 2.54 ($p = 0.011$) in the case of RE. Accounting for publication bias decreases the estimated $\gimel$ for fixed and random effects by 1 and 6 percentage points (pp), respectively.

Table 1: Simple meta-analysis, Eurozone studies

			95% CI			
Method	Effects	Estimate	Lower	Upper	z-stat.	Obs.
Standard	Fixed	0.038	0.033	0.043	14.917	28
Standard	Random	0.092	0.067	0.117	7.203	28
Trim and fill	Fixed	0.026	0.021	0.030	10.552	40
Trim and fill	Random	0.035	0.008	0.062	2.535	40
Impact weighted	Fixed	0.030	0.025	0.035	11.861	28
Impact weighted	Random	0.084	0.046	0.123	4.290	28

"Fixed effects": the true effect is assumed to be equal across all studies.

"Random effects": the true effect is assumed to be drawn from a normal distribution.

"Trim and fill": a nonparametric method by Duval & Tweedie (2000) accounting for publication bias.

Both estimates are decreased by another percentage point if studies are weighted according to the impact factor of the journal they were published in.[3] Moreover, applying jointly weights based on impact and number of citations (taken from RePEc) discounted by the article age, both fixed and random effects estimates become very small and insignificant even without correcting for publication bias. Robustness of this type of meta-analysis is usually checked by sensitivity analysis, when individual studies are gradually excluded from the sample and researchers investigate how the pooled estimate changes. In this case, there is one strong observation (Shin & Serlenga 2007), exclusion of which almost doubles the estimated ℷ for FE. However, the pooled estimate for RE is changed only by 0.4 pp. It is therefore safe to say that—according to this simple meta-analytical approach—the Rose effect of the Euro is lower than 6% (the upper bound of the CI corresponding to the RE estimate corrected for publication bias).

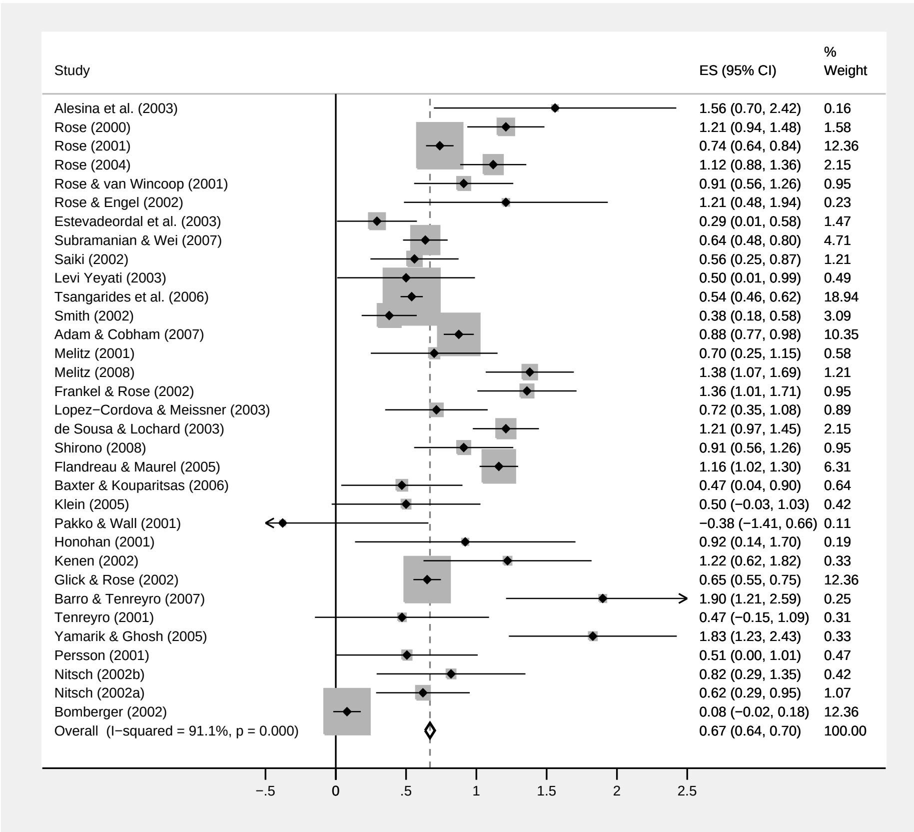

Figure 2: Forest plot of individual estimates, non-Euro studies

Forest plot of the results of non-Euro studies (Figure 2) shows a different picture. The pooled FE estimate is far from zero, namely 0.67 (ℷ = 95.42%),

[3]The process of assigning weights to working papers and other unpublished manuscripts is described in Table 13. However, the results would not change significantly if the weights for working papers were set to zero; neither do they alter when the Article Influence Score™ is used instead of the impact factor. It should be noted that using weights based on impact and citations is not an orthodox meta-analytical approach since it introduces *a priori* assumptions about the quality of research.

Table 2: Simple meta-analysis, non-Euro studies

Method	Effects	Estimate	95% CI Lower	95% CI Upper	z-stat.	Obs.
Standard	Fixed	0.670	0.636	0.704	38.116	33
Standard	Random	0.818	0.684	0.952	11.961	33
Trim and fill	Fixed	0.631	0.597	0.665	36.850	40
Trim and fill	Random	0.662	0.529	0.795	9.757	40
Impact weighted	Fixed	0.584	0.549	0.618	33.205	33
Impact weighted	Random	0.869	0.550	1.188	5.340	33

"Fixed effects": the true effect is assumed to be equal across all studies.

"Random effects": the true effect is assumed to be drawn from a normal distribution.

"Trim and fill": a nonparametric method by Duval & Tweedie (2000) accounting for publication bias.

CI = (88.89%, 102.18%); RE report even 0.818 ($\gimel$ = 126.6%), CI = (99.05%, 159.09%). Both estimates are highly significant with z-statistics 38.12 and 11.96, respectively. Again, it is evident that the confidence intervals often do not overlap, thus heterogeneity is probably present [Q-test: $\chi^2_{(32)} = 361$, $p < 0,001$] and RE are more reliable. Applying the trim and fill method to accomodate publication bias, the FE estimate changes only slightly, but the RE estimate drops significantly to 0.662 ($\gimel$ = 93.87%), CI = (69.72%, 121.44%)—see Table 2 for a comparison of different estimators in the case of non-Euro studies. Weighting by impact lowers the FE estimate of $\gimel$ by 16 pp; the RE estimate, however, is boosted by 12 pp. Adding citations weights almost doubles the estimates (1.351 for RE), but it also greatly increases the variance, even though both estimates remain highly significant. The most influential observation is Bomberger (2002); once it is excluded, the basic RE estimate slightly increases to 0.843 ($\gimel$ = 132.33%). Therefore, one might suspect the "true" Rose effect of currency unions (other than Eurozone) to be higher than what the RE trim and fill method would suggest. In any case, it is highly significant and seems to exceed 60% using this methodology. Assuming that currency unions double trade, as, e.g., Masson (2008) does when he asseses the welfare effects of forming currency unions in Africa, thus appears plausible in this respect.

There is no doubt that the estimates of the Rose effect of the Euro and other currency unions are indeed immensely different and that it is not very appropriate to pool them together. If one does so, heterogeneity is multiplied [Q-test: $\chi^2_{(60)} = 2071$, $p < 0,001$] and the difference between the simple FE and RE estimators increases; see Table 3. Worse still, correcting for publication bias and applying RE yields 0.074 ($\gimel$ = 7.68%), which is outside the Euro's 95% CI, but also very far away from the non-Euro estimate.

3 Publication Bias and the True Effect

In his thorough and influential review of the Rose effect literature, Richard Baldwin comments the meta-analysis of Rose & Stanley (2005): *"The meta-analysis statistical techniques are fascinating, but I don't believe it adds to our knowledge since deep down they are basically a weighted average of all point estimates."*

Table 3: Simple meta-analysis, all studies

Method	Effects	Estimate	95% CI Lower	95% CI Upper	z-stat.	Obs.
Standard	Fixed	0.050	0.045	0.055	20.179	61
Standard	Random	0.340	0.299	0.380	16.474	61
Trim and fill	Fixed	0.037	0.032	0.042	15.076	89
Trim and fill	Random	0.074	0.030	0.118	3.280	89
Impact weighted	Fixed	0.031	0.026	0.036	12.329	61
Impact weighted	Random	0.581	0.513	0.650	16.716	61

"Fixed effects": the true effect is assumed to be equal across all studies.

"Random effects": the true effect is assumed to be drawn from a normal distribution.

"Trim and fill": a nonparametric method by Duval & Tweedie (2000) accounting for publication bias.

(Baldwin 2006, p. 36). While this statement—or at least its last sentence—applies to the simple meta-analysis performed in Section 2, it disregards the most important part of Rose & Stanley (2005), as well as of the present study: the meta-regression analysis (MRA). The purpose of the meta-analytical methods performed Section 2 was to present a robustness check to the preferred MRA methodology since there are only 28 observations on the Eurozone at our disposal, which is slightly below the usually recommended threshold for regression analysis.

In this section, the MRA is employed to test for publication bias and the true underlying Rose effect. Publication selection can take the following two forms (Stanley 2005a):

Type I bias This form of publication bias occurs when editors, referees, or authors prefer a particular direction of results. Negative estimates of γ, for instance, might be disregarded; everybody knows that common currency does not hamper trade among the monetary union's members, right? The problem is that even if the true effect is positive, a certain percentage of studies (due to the laws of probability) should report negative numbers. Otherwise, the average taken from the literature can highly exaggerate the estimated true effect (TE). For instance, Stanley (2005a) shows that price elasticity of water demand is exaggerated fourfold due to publication bias.

Type II bias The second type of bias arises when statistically significant results are preferred; i.e., when editors choose "good stories" for publication. In this way, a completely fabulous effect may be "discovered" and further supported by subsequent research. Intra-industry spillovers from inward foreign direct investment might serve as an example (Görg & Greenaway 2004; Havránek & Iršová 2008).

The presence of type I publication bias is usually investigated employing the so-called funnel plot which shows the estimated effect against its precision (inverse of its standard error, Egger *et al.* 1997). The essence of this visual test is that, in the case of no bias, the shape of the cloud of observations must resemble an inverted funnel; observations with high precision should be concentrated

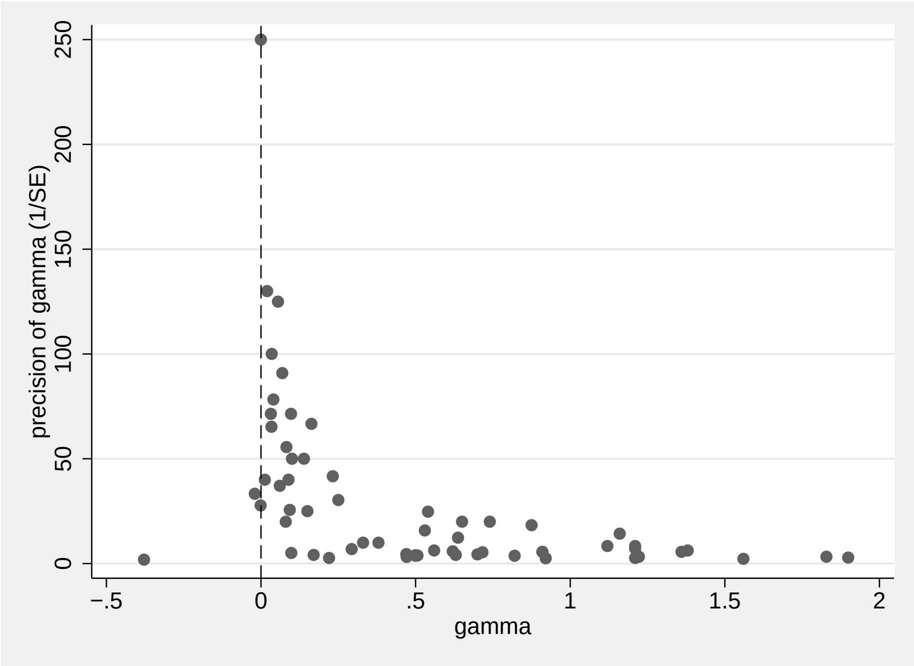

Figure 3: Funnel plot, all studies

closely to the TE, while those with lower precision should be more dispersed. In the absence of publication bias, the funnel must be also symmetric.

In Figure 3, the funnel plot for all 61 studies is presented. It shows a perfect example of strong publication bias. While positive estimates nicely form one half of a funnel, the left half is almost completely missing as there are only 4 non-positive estimates. The Eurozone and non-Euro studies taken separately resemble an inverted funnel even less. This test can be formalized using a simple MRA (Ashenfelter *et al.* 1999):

$$\gamma_i = \beta + \beta_0 SE_i + \mu_i, \quad i = 1, \ldots, M, \tag{2}$$

where M is the number of studies, β denotes the true effect and β_0 measures the magnitude of publication bias. However, regression (2) is evidently heteroskedastic. The measure of heteroskedasticity is the standard error of γ (SE), thus weighted least squares can be performed by running a simple OLS on equation (2) divided by SE:

$$t_i = \beta_0 + \beta \left(\frac{1}{SE_i} \right) + \vartheta_i. \tag{3}$$

The meta-response variable changes to t-statistic corresponding to γ_i. A simple t-test on the intercept of (3) is then a test for publication bias: funnel asymmetry test (FAT). But since the meta-explanatory variable ($\frac{1}{SE}$) contains errors (it was estimated), FAT is biased (Sterne *et al.* 2000). Fortunately, a remedy has been proposed: funnel-asymmetry heteroskedasticity-robust instrumental variables estimator (FAIVEHR, Davidson & MacKinnon 2004; Stanley

2005a) which uses the square root of the sample size as an instrument for $\frac{1}{SE}$ following Begg & Berlin (1988). Another approach is to use this square root directly as a proxy for $\frac{1}{SE}$ in (3).

Results of all three tests in the case of the Eurozone studies are summarized in Table 4. In all specifications, the intercept is highly significant (t-statistics vary from 2.91 to 6.02)—the hypothesis of no type I publication bias is thus strongly and robustly rejected, the rather that these tests are usually believed to have relatively low power (Stanley 2005a). The fact that they all reject the null hypothesis at the 99% level of significance implies that publication bias presents a very serious and widespread problem for the literature on the Euro's Rose effect.

Table 4: Tests of publication bias, Eurozone studies

	FAT-PET	FAIVEHR	PROXY
prec (effect)	0.000667	-0.0116	
	(0.05)	(-0.65)	
sqrtn (effect)			-0.000110
			(-0.67)
Constant (bias)	3.755^{**}	4.435^{**}	3.817^{**}
	(4.04)	(2.91)	(6.02)
Observations	28	28	28
RMSE	3.169	3.116	3.168

Huber-White heteroskedasticity-robust t-statistics in parentheses.
Meta-response variable: tstat.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

Type II bias can be assessed using the Galbraith plot (Galbraith 1988) that depicts the precision of γ against the t-statistic corresponding to the (assumed) TE. If the "true" effect was really true, only about 5% of the studies' t-statistics should exceed 2 in the absolute value. Figure 4 shows the Galbraith plot for the Eurozone studies (Galbraith plots for all or non-Euro studies yield similar results). If the TE was 0.05, 13 studies out of 28 would report significant results. The goodness of fit test easily rejects the hypothesis of the expected distribution [$\chi^2_{(1)} = 96$, $p < 0.001$]; the null hypothesis is rejected even more powerfuly when the TE is considered to be equal to 0 or 0.1. Therefore, type II bias is clearly present among the Eurozone studies.

All three methods of detecting type I bias (Table 4) can be also used to test for the significance and magnitude of the true effect corrected for publication bias [recall (2)]. Specifically, running a t-test on the slope coefficient of (3) is denoted as the precision effect test (PET). There are also other approaches; e.g., the meta-significance test (MST, Stanley 2001):

$$\log |t_i| = \phi_0 + \phi \log N + \kappa_i. \tag{4}$$

MST uses the statistical property that t-statistic should increase proportionally (in magnitude) to the square root of the number of observations. Significance and positiveness of ϕ is considered as an evidence for the existence of a TE beyond publication bias. MST, however, does not tell anything about the magnitude of the TE. Stanley (2006) developed the two-stage precision effect test

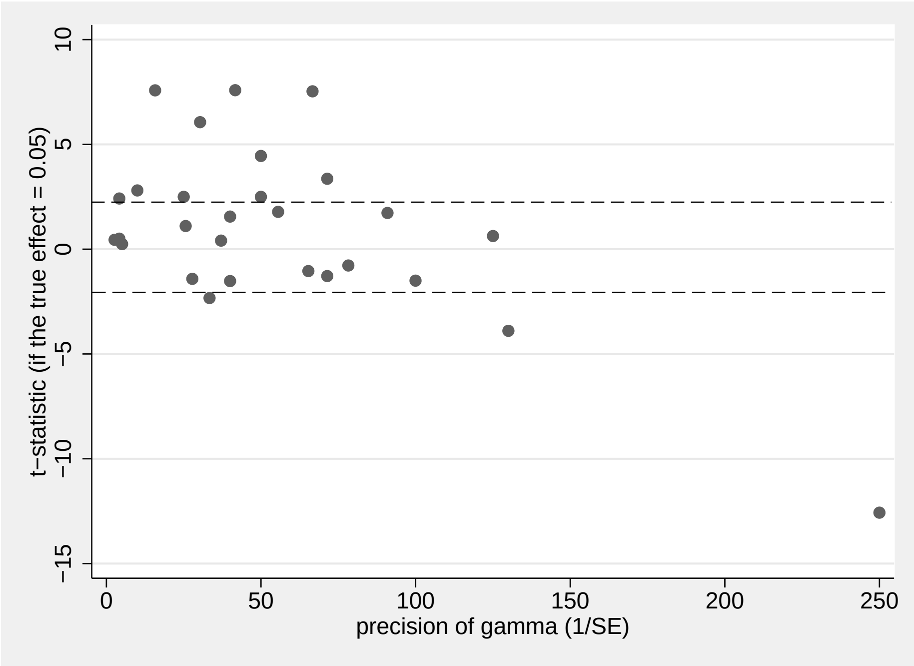

Figure 4: Galbraith plot, Eurozone studies

(PETS) that assumes publication bias to be quadratic in SE and uses corrected t-statistics (original t-stat. $-$ estimated bias) for (3) to estimate the TE more precisely. Another possibility is to modify (3) to account for type II bias:

$$|t_i| = \varphi_0 + \varphi \left(\frac{1}{SE_i} \right) + \omega_i. \tag{5}$$

The second stage is then similar to PETS.

Tests of the TE for the Eurozone studies are summarized in Table 5. Together with Table 4, there are six estimates of the TE corrected for publication bias. Five of them are completely insignificant (highest t-value: 0.05); three are even negative (FAIVEHR, PROXY, MST). Only PETS is positively significant at the 95% level. However, Stanley (2006) recommends using PETS only if both MST and PET first pass the test for effect, which is not the case here. Worse still, the quadratic form is not significant in the first stage of PETS, thus this method cannot be relied upon. When robust (iteratively re-weigthed least squares) versions of the tests of effect are used, the results do not change significantly. This means that there is not even a slight trace of any true underlying Rose effect of the Euro.[4]

Table 6 and Table 7 summarize the tests of publication bias and the TE for non-Euro studies. It is apparent that publication bias is weaker than in

[4]There is an obvious objection to this approach: if the Rose effect of the Euro is growing over time (Bun & Klaassen 2002; Baldwin 2006), how can one pool together studies written in 2002 when the Euro was infant, and studies written in 2008 when it was preparing to celebrate its tenth birthday? Simply because explanatory meta-regression does not find any relation between the γ of the Eurozone studies and time. Also Frankel (2008b) concludes that the Euro's effect has stabilised after a few starting years.

Table 5: Tests of the true effect, Eurozone

	MST	PETS	Corr. PET
logn (effect)	−0.00745		
	(−0.09)		
prec (effect)		0.0373*	0.000115
		(2.27)	(0.01)
Constant	0.855		
	(1.03)		
Observations	28	28	28
RMSE	1.453	4.003	3.110

Huber-White heteroskedasticity-robust t-statistics in parentheses.

Meta-response variable: logt for MST, tstat for PETS and corrected PET.

† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

Table 6: Tests of publication bias, non-Euro studies

	FAT-PET	FAIVEHR	PROXY
prec (effect)	0.534**	0.531**	
	(4.08)	(2.95)	
sqrtn (effect)			0.0113*
			(2.15)
Constant (bias)	1.712*	1.738	3.615**
	(2.21)	(1.35)	(3.29)
Observations	33	33	33
RMSE	3.234	3.135	4.237

Huber-White heteroskedasticity-robust t-statistics in parentheses.

Meta-response variable: tstat.

† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

the previous case; one of the three measures is even insignificant (FAIVEHR). However, as has been already mentioned, these tests of publication bias are known to have relatively low power, which further decreases in combination with the instrumental variables approach. It is therefore safe to say that there is publication bias among non-Euro studies, although weaker than among the Eurozone studies.

Table 7: Tests of the true effect, non-Euro studies

	MST	PETS	Corr. PET
logn (effect)	0.148^{*}		
	(2.74)		
prec (effect)		0.669^{**}	0.526^{**}
		(7.38)	(6.21)
Constant	0.0439		
	(0.09)		
Observations	33	33	33
RMSE	0.741	3.355	3.184

Huber-White heteroskedasticity-robust t-statistics in parentheses.
Meta-response variable: logt for MST, tstat for PETS and corrected PET.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

All six tests for effect are significant at least at the 95% level. PETS and corrected PET (5) are designed to deliver the most precise estimates under these circumstances. In this case, we prefer corrected PET, because the quadratic form in the first stage of PETS is again not significant. The corrected PET reports 0.526 ($\gimel = 69.22\%$), CI = (42.48%, 95.23%). There is a caveat, though: Stanley (2005b) uses Monte Carlo simulations to show that PET is reliable only if $\sigma_{\vartheta}^2 \leq 2$. Otherwise, the estimate might be exaggerated by misspecification biases. In this case, $H_0 : \sigma_{\vartheta}^2 \leq 2$ is rejected $[\chi^2_{(32)} = 162,\ p < 0,001]$, thus the TE is probably lower than the estimated 69.22%. It can be, however, safely concluded that other-than-Euro currency unions are expected to boost trade by about one half.

Tests for all 61 studies are reported in Table 8 and Table 9. Publication bias is present (average t-stat. = 6.07) and traces of the TE are weak—only PETS and MST are significant, but PETS is unreliable for the same reasons as before and the MST result is only qualitative (i.e., it tells that there might be some effect, but nothing about its magnitude). Once again it shows that it is very useful to split the sample into Euro and non-Euro studies.

Figure 5 represents the funnel plot of all studies corrected for publication bias [using (5); observations with corrected $|\gamma| > 1$ are cut from the figure]. Contrary to Figure 3, the present funnel plot is clearly symmetric—this is how the literature *should* look like.

4 Explanatory Meta-Regression

MRA can also be employed to determine possible dependencies of study results on its design. In fact, it has been the primary focus of most economic

Table 8: Tests of publication bias, all studies

	FAT-PET	FAIVEHR	PROXY
prec (effect)	-0.00735	0.0446	
	(-0.78)	(0.44)	
sqrtn (effect)			0.000513
			(0.57)
Constant (bias)	5.134^{**}	3.596	4.807^{**}
	(8.02)	(1.20)	(9.00)
Observations	61	61	61
RMSE	4.112	4.593	4.115

Huber-White heteroskedasticity-robust t-statistics in parentheses.
Meta-response variable: tstat.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

Table 9: Tests of the true effect, all studies

	MST	PETS	Corr. PET
logn (effect)	0.118^{*}		
	(2.11)		
prec (effect)		0.0493^{*}	-0.00790
		(2.38)	(-1.03)
Constant	0.0648		
	(0.12)		
Observations	61	61	61
RMSE	1.158	5.664	4.078

Huber-White heteroskedasticity-robust t-statistics in parentheses.
Meta-response variable: logt for MST, tstat for PETS and corrected PET.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

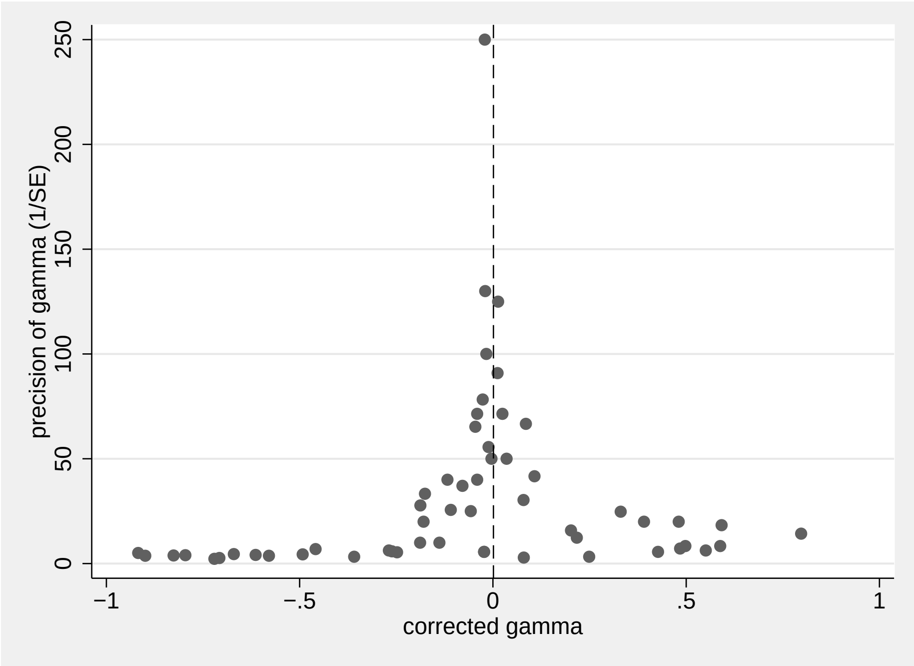

Figure 5: Funnel plot corrected for publication bias, all studies

meta-analyses since the pioneering work of Stanley & Jarrell (1989). Economics research is usually much more heterogeneous than epidemiology and psychology, where the meta-analytical approach was originally developed. In this respect, MRA is used to assign a pattern to heterogeneity.

We gathered 20 meta-explanatory variables that reflect study design, social attributes, and skills of the authors (see Table 13), 8 of them are assumed to affect publication bias, the rest 12 are expected to influence the estimates of γ directly. The former include researchers' gender, nationality, ranking, proxies of IT skills, panel nature of the data, and year of publication (for a list of possible regressors affecting publication bias, see Stanley *et al.* 2008). The latter cover dummies for specific authors, short run nature of the study, Eurozone data, postwar data, number of countries and years in the dataset, and impact factor. All meta-explanatory variables were chosen *ex ante*.

The simplest estimator used here is multiple MST—the same model as in (4) plus all additional meta-explanatory variables. Contrary to the previous sections, now the focus rests on the whole sample because more degrees of freedom are needed; heterogeneity is not so much problematic since it can be modeled to a large extent. There are 61 observations, which is enough for a meta-regression since sample size in meta-analysis is substantially more effective in increasing the power of hypothesis testing than sample size of original studies (Koetse *et al.* 2007). The most insignificant meta-regressors are excluded one by one to get a model which contains only variables significant at least at the 90% level (column 1 in Table 10). Heteroskedasticity-robust standard errors are used. The model passess Ramsey's RESET test [$F_{(3,50)} = 1.33$, $p > 0.05$] and there is no sign of multicolinearity (condition number = 18); however, the skewness-kurtosis test

Table 10: Explanatory meta-regression analysis: multiple MST

	all studies		non-Euro	Eurozone
	OLS	IRLS	OLS	OLS
logn	0.225^{**}	0.149^{*}	-0.0258	0.192^{*}
	(4.19)	(2.35)	(-0.30)	(2.52)
rose	0.744^{**}		0.491^{*}	
	(3.23)		(2.08)	
baldwin	-1.993^{**}	-1.139^{**}		-4.947^{**}
	(-7.16)	(-2.93)		(-3.26)
taglioni	1.816^{**}			1.696^{*}
	(3.94)			(2.41)
euro	-1.959^{**}	-0.815^{**}		
	(-3.47)	(-2.91)		
shortrun	1.571^{*}	1.394^{**}	0.995^{*}	$1.686^{\dagger}$
	(2.58)	(5.04)	(2.59)	(2.07)
countries	-0.005^{*}	-0.006^{**}		-0.012^{**}
	(-2.33)	(-3.09)		(-3.68)
repec		0.828^{**}		
		(3.80)		
latex		-1.205^{**}	$-0.484^{\dagger}$	
		(-4.05)	(-1.99)	
denardis		-0.958^{*}		
		(-2.11)		
tenreyro		1.291^{*}		
		(2.35)		
panel		0.904^{**}	1.130^{**}	
		(3.79)	(3.40)	
years		0.019^{**}	0.0293^{**}	$-0.044^{\dagger}$
		(2.86)	(3.28)	(-1.98)
nitsch			0.603^{*}	
			(2.06)	
topfive				2.841^{*}
				(2.23)
Constant	-0.148	-1.014^{*}	-0.143	-0.821
	(-0.40)	(-2.03)	(-0.23)	(-0.53)
Observations	61	61	33	28
R^2	0.413	0.601	0.590	0.633
Adjusted R^2	0.335	0.511	0.475	0.505

t-statistics in parentheses (Huber-White heteroskedasticity-robust for OLS).
Meta-response variable: logt.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

of normality rejects the null hypothesis [$\chi^2_{(2)} = 15.92$, $p < 0.001$] and also the fit of the model is not very good ($R^2 = 0.41$).

The fit can be improved by means of downweighting influential observations; iteratively re-weighted least squares (IRLS) present a robust alternative to OLS (Hamilton 2006, pp. 239–256). Robust methods in meta-analysis are employed, e.g., by Havránek & Iršová (2008). The respective model is summarized in column 2 of Table 10, R^2 increases to 0.60, and also adjusted R^2 is quite high considering the nature of the data (0.51). Both OLS and IRLS predict that studies with more cross-sectional units, studies on the Eurozone, and papers co-authored by Baldwin tend to report a lower Rose effect *ceteris paribus*; short run studies, on the other hand, report marginally higher effects. OLS additionally find Rose's and Taglioni's co-authorship significantly positive. IRLS report marginally negative effect of de Nardis's co-authorship and positive effect of Tenreyro's co-authorship, as well as positively significant coefficients for panel data and the number of years in the dataset. IRLS also show interesting results concerning the proxies for IT skills: authors registered in RePEc tend to report higher estimates, authors using LaTeX are expected to publish lower Rose effects.

Dividing the sample into the Eurozone and non-Euro studies and running simple OLS yields columns 3 and 4 in Table 10. The fit is considerably better (R^2 is 0.590 and 0.633, respectively) and both models pass the battery of specification tests (no multicollinearity, Ramsey's RESET test not significant, normality not rejected). What is new is the significance of Nitsch's co-authorship for non-Euro studies (positive effect) and ranking for Eurozone—if there is a top economist among co-authors, the study reports significantly higher effects. It is also interesting that the effect of the number of years in the dataset is opposite for Eurozone and non-Euro studies: negative and positive, respectively.

In the multiple MST framework, there was no distinction between variables that affect publication bias and regressors that influence γ, neither were the knowledge of the standard errors of γ utilized. Stanley *et al.* (2008) augment FAT-PET to the following multivariate version:

$$t_i = \underbrace{\beta_0 + \sum_{j=1}^{J} \theta_j S_{ji}}_{\text{bias}} + \underbrace{\tilde{\beta}}_{\text{pseudo TE}} \left(\frac{1}{SE_i}\right) + \underbrace{\sum_{k=1}^{K} \frac{\delta_k Z_{ki}}{SE_i}}_{\text{controls}} + \vartheta_i, \qquad (6)$$

where S_j is a set of variables influencing publication bias and Z_k is a set of variables affecting γ directly. Stanley *et al.* (2008) denote $\tilde{\beta}$ as the true effect, but it should rather be called "true effect of the baseline case" or pseudo TE (its magnitude, of course, depends on the units used and the definition of dummies and other regressors). We denote this estimator as fixed effects, even though it is not the traditional FE estimator used in meta-analysis: note that variables S_j are not divided by the standard error.

FE estimates are summarized in Table 11. After insignificant variables were excluded, the "economics research cycle hypothesis" (there is a predictable pattern of novelty and fashion in economics; initial path-breaking results are confirmed by other high estimates, but as the time passes, skeptical results become preferable: Goldfarb 1995; Stanley *et al.* 2008) was tested by adding the year of

Table 11: Explanatory meta-regression analysis: fixed effects

	all studies		non-Euro	Eurozone
	OLS	IRLS	OLS	OLS
prec	0.770**	0.565**	0.626*	0.0947**
	(5.30)	(3.98)	(2.21)	(4.17)
panel	1.414†	2.872**	4.018*	
	(1.74)	(3.21)	(2.76)	
rose_se	0.484**	0.462**		
	(3.23)	(3.90)		
nitsch_se	−0.151**	−0.252*		
	(−4.19)	(−2.52)		
baldwin_se	−0.0758**			−0.0764**
	(−5.17)			(−4.74)
denardis_se	−0.0395†			−0.0626**
	(−1.96)			(−4.61)
euro_se	−0.700**	−0.660**		
	(−5.30)	(−5.19)		
shortrun_se	0.0372*	0.115**		
	(2.03)	(3.27)		
countries_se	−0.0023**	−0.0025**	−0.006**	
	(−2.87)	(−3.44)	(−3.24)	
impact_se	−0.0432*			−0.0524**
	(−2.17)			(−2.93)
year	1.083	1.847*		
	(1.63)	(2.65)		
year2	−0.0669	−0.143†		
	(−0.90)	(−1.89)		
repec		1.708†		
		(1.89)		
topfive		−1.872*		
		(−2.27)		
latex		−2.713*	−3.322*	
		(−2.09)	(−2.59)	
tenreyro_se		1.338†	1.257†	
		(1.72)	(1.87)	
years_se		0.0029*	0.011†	−0.0009†
		(2.49)	(1.87)	(−1.74)
postwar_se			0.552†	
			(1.78)	
Constant	−1.307	−3.581†	−1.696	2.105*
	(−1.00)	(−1.97)	(−1.23)	(2.36)
Observations	61	61	33	28
R^2	0.712	0.713	0.734	0.496
RMSE	2.454	2.507	2.682	2.445

t-statistics in parentheses (Huber-White heteroskedasticity-robust for OLS).
Meta-response variable: tstat.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

publication and its square value.[5] The null hypothesis corresponds to the joint significance of these variables and concave shape of the relationship. In this case, $F_{(2,48)} = 3.85$, $p < 0.05$ and the relationship is indeed concave, hence the economics research cycle hypothesis is supported for this type of literature. This becomes even more apparent when IRLS are used [$F_{(2,46)} = 7.28$, $p < 0.01$]. On the other hand, the research cycle hypothesis is rejected when both groups of literature are considered separately: $F_{(2,23)} = 1.56$, $p > 0.05$ for non-Euro studies and $F_{(2,20)} = 0.21$, $p > 0.05$ for the Eurozone studies; there is no apparent dependence on time. This might suggest that the research cycle identified in the whole literature emerges also due to a higher proportion of the Eurozone papers in a few recent years.

Regression described in column 1 of Table 11 is not very well specified, however. Condition number is high (75), Ramsey's RESET is significant [$F_{(3,45)} = 4.42$, $p < 0.05$], only normality is not rejected [$\chi^2_{(2)} = 1.36$, $p > 0.05$].[6] FE bring much better fit compared to multiple MST: $R^2 = 0.71$ for both OLS and IRLS. The estimates remain similar to MST with a few exceptions: Nitsch's co-authorship has marginally negative effect for all studies (in MST, the effect was significant only for non-Euro studies and it was positive). Studies published in journals with high impact factor are more probable to report lower effects, as do articles co-authored by top economists (compared to marginally positive effect of the latter for the Eurozone studies in MST). This relationship is interesting, but it is not confirmed by other specifications. Additionally, postwar data were found to deliver significantly higher estimates for non-Euro studies. Compared to MST, Taglioni's co-authorship is not significant using fixed effects.

It is apparent that FE were able to model a significant portion of the heterogeneity inside the sample (recall high R^2s). Nevertheless, a lot of heterogeneity remain unexplained. Testing $H_0 : \sigma^2_{\vartheta} = 1$ (FE MRA explains heterogeneity well) yields $\chi^2_{(60)} = 289$, $p < 0.001$; for column 1, therefore, H_0 is rejected (the result is qualitatively the same also for the other specifications). When this is the case, random effects MRA might be preferable (see, e.g., Abreu *et al.* 2005):[7]

$$\gamma_i = \iota_0 + \sum_{j=1}^{J} \theta_j S_{ji} + \sum_{k=1}^{K} \delta_k Z_{ki} + \lambda_i + \rho_i, \tag{7}$$

where λ_i stands for a normal disturbance term with standard deviations assumed to be equal to SE_i, and ρ_i is a normal disturbance term with unknown variance τ^2 assumed equal across all studies. This between-study variance is estimated using the residual maximum likelihood method; t-values are computed employing the Knapp & Hartung (2003) modification.

RE MRA is summarized in Table 12. There are much less significant explanatory variables. For this reason, in columns 2, 3, and 4, all regressors with p-values lower than 0.2 were retained in the model. The most interesting result

[5] This test is meaningful when t-statistic is the meta-response variable—that is why it is not applied in the case of multiple MST or random effects MRA.

[6] The model passes all specification tests if variables *panel*, *year*, and *year2* are excluded. Specifications in columns 3 and 4 pass the battery of specification tests right away.

[7] Monte Carlo experiments suggest that RE MRA is preferable if heterogeneity is caused by variance in γ or differences in the true underlying effect across studies. When heterogeneity arises due to omitted variable bias, however, the other estimators should be relied upon (Koetse *et al.* 2007). For this reason, FE MRA and multiple MST are interpreted here as well, even though more weight is given to random effects.

Table 12: Explanatory meta-regression analysis: random effects

	all studies		non-Euro	Eurozone
	(1)	(2)	(3)	(4)
woman	0.206*	0.169†	0.499**	
	(2.32)	(1.89)	(2.82)	
rose	0.295*	0.284*		
	(2.33)	(2.26)		
baldwin	−0.274†	−0.283†		−0.667**
	(−1.89)	(−1.98)		(−7.27)
euro	−0.593**	−0.608**		
	(−7.38)	(−6.13)		
repec		0.132		
		(1.53)		
denardis		−0.230		−0.0810*
		(−1.46)		(−2.54)
tenreyro		0.514†		
		(1.78)		
shortrun		0.160		0.0856†
		(1.59)		(1.87)
topfive			0.386*	0.182*
			(2.48)	(2.74)
nitsch			0.458	
			(1.33)	
panel			0.447	
			(1.48)	
logn				0.0457**
				(5.42)
usa				−0.0551
				(−1.62)
taglioni				0.397**
				(5.74)
years				−0.00210
				(−1.43)
Constant	0.679**	0.554**	0.0562	−0.297*
	(10.01)	(5.87)	(0.18)	(−2.63)
Observations	61	61	33	28
τ^2	0.0473	0.0423	0.1024	0.0014

t-statistics in parentheses, Knapp & Hartung (2003) modification used.
Method for estimating between-study variance: residual maximum likelihood.
Meta-response variable: gamma.
† $p < 0.10$, * $p < 0.05$, ** $p < 0.01$

is the significance of authors' gender: if there is at least one woman among co-authors, the study is probable to find higher Rose effect. This pattern is particularly strong for non-Euro studies. If all co-authors of a paper are Americans and study the trade effect of the Euro, they are likely to report lower results.

It is clear from the conducted tests that explanatory meta-regression is as sensitive to method and specification changes as any other field of empirical research. The most important meta-explanatory variables are those that can be found in more specifications, especially random effects, and that retain their signs (effect on γ in parentheses): studies on the Eurozone $(-)$, Rose's co-authorship $(+)$, Baldwin's co-authorship $(-)$, Tenreyro's co-authorship $(+)$, de Nardis's co-authorship $(-)$, authors registered in RePEc $(+)$, short run nature $(+)$, manuscript written in LaTeX$(-)$, number of countries in the dataset $(-)$, and panel data $(+)$.

5 Conclusion

Literature on the trade effect of currency unions is heterogeneous to a large extent. Studies estimating the trade effect of the Euro find on average much smaller effects than articles concentrating on other currency unions. The present meta-analysis shows that it is not appropriate to pool all these estimates together in a search for the "true effect."

Evidence for publication selection—i.e., preference towards statistically significant and positively biased results—is robust among both Eurozone and non-Euro studies, although it is much stronger for the former. Narrative literature reviews that do not take publication selection into account (e.g., Frankel 2008b) are hence vulnerable to a substantial upward bias. Meta-regression methods show that, beyond publication bias, there is a significant and huge Rose effect of the other-than-Euro currency unions, safely about 50% and maybe more; but no effect at all for the Euro. Using a simpler approach of correcting for publication bias (the trim and fill method), the estimated Rose effect of the Euro becomes significant, but not economically important (3.56%), and weighting by the impact factor and citations further decreases the estimate. The absence of an economically important true effect is so robust that even some possible mistakes in the process of chosing the authors' preferred estimates cannot easily change the outcome.

Employing explanatory meta-regression, about 70% of the heterogeneity among the "Rosean" literature can be modeled. Unfortunately, the authorship of a particular study is especially important: papers co-authored by Rose or Tenreyro tend to find higher effects, papers co-authored by Baldwin or de Nardis are more likely to report smaller estimates. Papers on the Eurozone find significantly lower effects as well as long run studies and studies with a high number of countries in their datasets do. When all co-authors are registered in RePEc, the study tends to report higher effects. Once a draft of the study is written in LaTeX, it is probable to find a rather smaller Rose effect. The Rose effect literature taken as a whole shows signs of the economics research cycle (Goldfarb 1995; Stanley *et al.* 2008): reported t-statistic is a quadratic function of publication year. One might take a note that the literature seems to have almost completed the circle and the results, especially on the Eurozone, are getting close to those "before Rose" when exchange rate volatility was believed to have low influence on international trade (McKenzie 1999).

The present author does not dare to argue that the Euro would have no effect on trade. The effects may indeed vary from country to country and industry to industry, as Baldwin (2006) suggests. At the very least, however, there is something wrong with the present Rosean literature applied on the Eurozone. The degree of publication bias is striking and the trade effect of the Euro is probably much lower than we believed, even if "what we believed" was already twentyfold less than what Rose reported in his famous article.

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A Suplementary Tables

On the following pages, a few additional illustrative tables are provided.

Table 14: Studies used in the meta-analysis

Study	**Euro**	**Gamma**	***t*-stat.**	**Impact**
Rose (2000)	no	1,2100	8,643	1,281
Pakko & Wall (2001)	no	-0,3780	-0,715	0,536
Rose & van Wincoop (2001)	no	0,9100	5,056	2,239
Rose (2001)	no	0,7400	14,800	1,281
Persson (2001)	no	0,5060	1,969	1,281
Honohan (2001)	no	0,9210	2,303	1,281
Mélitz (2001)	no	0,7000	3,043	0,036
Tenreyro (2001)	no	0,4710	1,491	0,018
Nitsch (2002b)	no	0,8200	3,037	0,715
Frankel & Rose (2002)	no	1,3600	7,556	3,688
Thom & Walsh (2002)	yes	0,0980	0,500	0,994
Glick & Rose (2002)	no	0,6500	13,000	0,994
Rose & Engel (2002)	no	1,2100	3,270	0,947
Bun & Klaassen (2002)	yes	0,3300	3,300	0,018
de Souza (2002)	yes	0,1700	0,708	0,018
Nitsch (2002a)	no	0,6200	3,647	0,018
Smith (2002)	no	0,3800	3,800	0,018
Bomberger (2002)	no	0,0800	1,600	0,018
Saiki (2002)	no	0,5600	3,500	0,018
Kenen (2002)	no	1,2219	4,006	0,018
Levi Yeyati (2003)	no	0,5000	2,000	0,302
Estevadeordal *et al.* (2003)	no	0,2930	2,021	3,688
Barr *et al.* (2003)	yes	0,2500	7,576	1,281
Lopéz-Córdova & Meissner (2003)	no	0,7160	3,849	2,239
Micco *et al.* (2003)	yes	0,0890	3,560	1,281
de Nardis & Vicarelli (2003)	yes	0,0610	2,262	0,018
Cabasson (2003)	yes	0,6300	2,625	0,018
Alesina *et al.* (2003)	no	1,5600	3,545	0,036
de Sousa & Lochard (2003)	no	1,2100	10,083	0,018
Rose (2004)	no	1,1200	9,333	2,239
de Nardis (2004)	yes	0,0930	2,385	0,382
Sadikov *et al.* (2004)	yes	0,2200	0,579	0,036
Faruqee (2004)	yes	0,0820	4,556	0,036
Taglioni (2004)	yes	0,5300	8,370	0,018
Baldwin & Taglioni (2004)	yes	0,0340	2,220	0,018
Flandreau & Maurel (2005)	no	1,1600	16,571	0,143
Klein (2005)	no	0,5000	1,852	0,709
Yamarik & Ghosh (2005)	no	1,8285	6,000	0,072
Aristotelous (2006)	yes	0,0550	6,875	0,653
Flam & Nordström (2006a)	yes	0,2320	9,667	0,036
Baldwin & Taglioni (2006)	yes	-0,0200	-0,667	0,036
Baldwin & di Nino (2006)	yes	0,0350	3,500	0,036

Continued on next page

Table 14: Studies used in the meta-analysis (continued)

Study	Euro	Gamma	*t*-stat.	Impact
Flam & Nordström (2006b)	yes	0,1390	6,950	0,018
Gomes *et al.* (2006)	yes	0,0690	6,273	0,018
Tsangarides *et al.* (2006)	no	0,5400	13,370	0,036
Baxter & Kouparitsas (2006)	no	0,4700	2,136	0,036
Barro & Tenreyro (2007)	no	1,8990	5,410	0,535
Subramanian & Wei (2007)	no	0,6370	7,864	1,541
Adam & Cobham (2007)	no	0,8750	16,010	0,153
Shin & Serlenga (2007)	yes	-0,0003	-0,075	1,094
Bun & Klaassen (2007)	yes	0,0320	2,286	0,732
de Sousa & Lochard (2007)	yes	0,1500	3,750	0,018
Shirono (2008)	no	0,9100	5,056	0,072
Mélitz (2008)	no	1,3800	8,625	0,994
Berger & Nitsch (2008)	yes	-0,0010	-0,028	0,709
Brouwer *et al.* (2008)	yes	0,0120	0,480	0,709
Baldwin *et al.* (2008)	yes	0,0200	2,600	0,036
Cafiso (2008)	yes	0,1630	10,867	0,036
de Nardis *et al.* (2008)	yes	0,0400	3,130	0,072
Frankel (2008b)	yes	0,0970	6,929	0,036
Chintrakarn (2008)	yes	0,1000	5,000	0,072

Impact factor for the year 2007 obtained from ISI Web of Knowledge (see Table 13).

Table 13: Acronyms of regression variables

Variable	Explanation
gamma	Point estimate of common currency's effect on trade.
tstat	t-statistic corresponding to gamma.
logt	Natural logarithm of absolute value of *tstat*.
se	Standard error of *gamma*.
prec	Inverse of *se*.
sqrtn	Square root of number of observations.
logn	Natural logarithm of number of observations.
	Moderator variables affecting publication bias
woman	= 1 if there is a woman among co-authors, zero otherwise.
usa	= 1 if all co-authors are Americans (based on current address).
repec	= 1 if all co-authors are registered in RePEc.
topfive	= 1 if at least one co-author ranks among top 5% in at least 10 cathegories on RePEc.
latex	= 1 if an earlier draft of the study or the study itself (in case it is unpublished) was written in LaTeX.
panel	= 1 if the study uses panel data with N > T.
year	Publication year−2000.
year2	Variable *year* squared.
	Moderator variables affecting gamma directly
rose	= 1 if Rose is a co-author.
nitsch	= 1 if Nitsch is a co-author.
baldwin	= 1 if Baldwin is a co-author.
denardis	= 1 if de Nardis is a co-author.
taglioni	= 1 if Taglioni is a co-author.
tenreyro	= 1 if Tenreyro is a co-author.
euro	= 1 if the study concentrates on the Eurozone.
shortrun	= 1 if the study has short run character.
countries	Number of countries in the dataset.
years	Number of years in the dataset.
postwar	= 1 if postwar data are used.
impact	Impact factor of the journal where the study was published. Journals without an impact factor obtain weights corresponding to 50% of the lowest impact factor in this sample. Working papers by NBER, ECB, European Commission, CESifo, and CEPR obtain 25%. Other unpublished manuscripts get 12.5%.

Variables divided by *gamma*'s standard error are denoted as *variable*_se (e.g., Table 11).

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