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Abstract

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JEL Classification: F14, G21, G28

Keywords: international trade finance, Basel II, letters of credit

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Risk-Based Capital Requirements for Banks and International Trade *

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October 7, 2016

Abstract

We find that changes in banks' risk-based capital requirements can affect firm-level exports. We exploit the mandatory Basel II adoption in its Standardized Approach by all banks in Turkey on July 1, 2012. This change affects risk-weights for letters of credit and generates two identification schemes with opposite predicted signs. Using data that cover 16,662 exporters shipping 2,888 different products to 158 countries, we find that the share of letter of credit-based exports decreases (increases) at the firm-country-product level when the associated counterparty risk-weights increase (decrease) after Basel II adoption. However, growth of firm-product-country level exports remains unaffected. (98 words)

JEL: G21, G28, F14

Keywords: letters of credit; international trade finance; exports; risk-weights; Basel II

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Risk-Based Capital Requirements for Banks and International Trade

1. Introduction

Many institutions involved in global trade raise serious concerns regarding the treatment of financial instruments related with international trade under later versions of risk-based capital requirements proposed by the Basel Committee on Bank Regulation of the Bank for International Settlements (BIS). For example, in 2009 Robert Zoellick, the then president of the World Bank, suggested that 10%-15% of the decrease in global trade during the Great Recession might be due to lower provision of trade finance under Basel II (Financial Times, February 19, 2009).[1] A 2009 survey by the International Chamber of Commerce (ICC) reports that "the feedback … on Basel II … [suggests] that most banks are facing tougher capital requirements for their [international] trade assets" (ICC, March 31, 2009, p. 40). Other banking surveys indicate that (i) Basel II had a negative impact on banks' provision of trade finance for most large international financial institutions and that (ii) for a non-negligible proportion of banks the increase in the cost of trade finance products is linked with higher capital requirements (Asmundson et al., 2011). Given such worries, during its Seoul Summit the G20 stated that it would "… evaluate impact of regulatory regimes on trade finance" (G20, 2010).

To shed some light on these concerns, we examine the impact of Basel II adoption on the letter of credit-financed exports.[2] Letters of credit are standard international trade financing instruments that are issued by the importer's bank.[3] The importer would then send the instrument to the exporter. The latter would present the letter of credit it receives to its local bank, together with other documents needed as proof of export transaction, for payment. The letter of credit issuing and receiving banks have to hold it as an off-balance sheet item, which creates a capital charge for both institutions.[4] For the bank that holds the letter of credit, the related capital requirement is calculated by multiplying the nominal value of the letter of credit first by a *credit conversion factor* to obtain the on-balance sheet equivalent, and then, with a *risk-weight* to adjust for counterparty-bank risk exposure. In this paper we focus on the specific case of a country whose particular form of Basel II adoption on July 1st 2012 allows us to isolate the effects of changes in capital requirements on exports. This is because the Turkish banking regulators required all banks operating in the country to move from Basel I to the Standardized Approach (SA)

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[1] Similar fears have been raised for Basel III, which initially proposed that a 100% capital earmark be set aside for many off-balance sheet items, including letters of credit (see for example, Financial Times, October 19, 2010 and the Wall Street Journal, February 6, 2011). Upon consultations with the World Bank, the World Trade Organization (WTO), and the ICC, the BIS relaxed certain aspects of capital requirements for international trade instruments under Basel II and III (Financial Times, October 25, 2011, and BIS, October 2011).

[2] These are commercial letters of credit, rather than standby letters of credit that are typically used for credit enhancement.

[3] Other main methods of payment in international trade are "cash in advance" (in which the importer bears the transaction risk by paying the exporter prior to shipment), and "open account" (in which the exporter bears the transaction risk by getting paid by the importer after the reception of goods).

[4] In some countries, for example, the US, the confirmed-letters of credit can be sold in the money market as bankers' acceptances. In the case that we study, Turkey, there is no such secondary market: once confirmed, letters of credit have to be held as off-balance items until maturity.

version of Basel II, whose components are public knowledge.[5] This resulted in changes in risk-weights while the credit conversion factors remained constant.[6] Under Basel II risk-weights either increase or decrease depending on the letter of credit-issuing counterparty bank's agency rating, whereas under Basel I they were fixed depending on whether the counterparty bank is located in an OECD-member country or not. This overlay of Basel I and II rules results in two sets of identification schemes in which the impact of risk-weight changes is expected go in opposite directions for exports to OECD and non-OECD countries.

More specifically, we examine how Turkey's adoption of Basel II on July 1st 2012 affected its firms' letter of credit-based share of exports at the firm-country-product level. Under Basel I, for a counterparty bank that is located in an OECD-member country, the risk-weight is 20%. For counterparties located in non-OECD countries the risk-weight is 100% prior to July 1, 2012. In contrast, under the Standardized Approach version of Basel II after July 1, 2012 the risk-weights change based on (i) the remaining maturity of the letters of credit (shorter or longer than three months), and (ii) the national regulator-defined groups of agency-rating categories following Bank for International Settlements (BIS) guidelines. These changes in risk-weights, when combined with the simpler standard under Basel I, generate testable implications with opposite signs that differ based on whether Turkish exports are destined to an OECD-member country or not.

For counterparty banks located in OECD countries and for letters of credit with remaining maturities longer than three months, the associated risk-weights either (i) increase by 150% (from 20% under Basel I to 50% under Basel II) for lower investment grade (A1 to Baa3 by Moody's, or equivalently A+ to BBB- by S&P or Fitch) rated or non-rated counterparty banks, or (ii) stay constant at 20% for higher investment grade (Aaa to Aa3, or equivalently AAA to AA-) rated counterparty-banks. The latter category forms the base-case in our difference-in-differences regressions.[7] In this particular case, we hypothesize that Turkish export shares involving letters of credit issued by A1 to Baa3 rated counterparty banks would decrease due to pricing and/or credit-exposure limit effects. The pricing channel would arise if Turkish banks pass on the higher cost of capital to their exporter clients: as the associated risk-weight increases, holding an export-letter of credit from A1 to Baa3 rated counterparty banks becomes more expensive compared to Aaa to Aa3 rated counterparties. An

[5] In contrast, during Basel II adoption in other countries, the banking regulators typically require large and/or sophisticated banks to adopt Internal Ratings Based Approach (IRB) under its Foundations or Advanced versions. IRBs, and hence their components, are proprietary to the institutions that adopt them, making identification difficult, if not impossible, for tests such as ours.

[6] Under Basel I and II, the credit conversion factors can differ between 20%, 50% or 100% across the types of binding letters of credit that require a payment. But they remain fixed per letter of credit category during period we study. See Section 4.1 for more details.

[7] In fact, three additional categories are also possible post-Basel II for letters of credit with maturities higher than three months that are issued by OECD banks: risk-weights (i) increased to 100% for non-investment grade and non-default rated counterparties, and (ii) went up to 150% for imminent or actual default categories (i.e., Caa1 or CCC+ and below rated counterparties). But these cases do not, in effect, apply in our particular setting because of the restrictions we need to impose on the data for a proper difference-in-differences estimation.

alternative channel would arise if Turkish banks adopt credit-exposure limits as a simple risk-management tool and ration holding letters of credit for certain rating class ranges (in this case to limit their overall exposure to A1 to Baa3 rated counterparties). These two channels are not mutually exclusive, but we have no way of differentiating among them.

For exports to non-OECD countries a second set of identification schemes apply. Suppose, for the sake of an example, that a Turkish exporter presents its Turkish bank a letter of credit issued by a counterparty bank that is located in a non-OECD country. The letter of credit is for $ 1 million (approximately equal to 1.8 million Turkish Liras [TL] on July 2, 2012) and has *longer* than three months of remaining maturity. Prior to July 1, 2012, under Basel I, given that the counterparty bank is located in a non-OECD country, holding this export letter of credit would have required that the Turkish bank sets aside $ 120,000 in additional capital, *irrespective* of the risk of the counterparty-bank.[8] Under Basel II, the capital charge would either (i) decline by 80% to $ 24,000 if the letter of credit-issuing counterparty bank is rated Aaa through Aa3, (ii) drop by 50% if the importer's bank has a rating between A1 through Baa3 or is not rated, (iii) remain equal to $ 120,000 if the counterparty bank's rating is between Ba1 and B3, or (iv) would increase by 50% to $ 180,000 if the counterparty-bank is rated Caa1 and below. The underlying mechanisms remain the same as the one for the OECD-based letter of credit-issuing counterparties.[9]

To test these conjectures, we use difference-in-differences models, estimated using Ordinary Least Squares (OLS), where the dependent variable is the letter of credit-based share of exports at the firm-country-product-level. Using the share of letter of credit-based exports allows us to implicitly control for demand effects. We take the first-difference of the share of exports in order to account for the observed differences in pre-Basel II time-trends across the treatment groups (for which the risk-weights change) and the control groups (for which the risk-weights remain the same). This triple-difference approach allows us to control for confounding factors at the firm-country-product level so long as they remain constant over two consecutive annual periods.[10] We also note that for any unaccounted variation to influence our findings, the remaining confounding factors' impact would need to be systematic enough to generate the same effects (with the opposite signs) in two separate (OECD and non-OECD) samples. This is an economically implausible scenario given that our firm-country-product triplets cover 16,662 exporters, 158 destinations, and 2,888 six-digit Harmonized System (HS6) product categories.

Our findings support the conjectures of the previous paragraph. For the OECD sample, our base-line regressions indicate that the share of letter of credit-based exports decrease by 11.80% given the

[8] $ 120,000 = $ 1,000,000×1.00×1.00×0.12, where the credit conversion factor is equal to 100% (for a confirmed export-letter of credit), risk-weight is equal to 100% (for a non-OECD counterparty under Basel I), and the minimum Tier 1+Tier 2 capital ratio is equal to 12% (as required by Turkish banking regulators).

[9] For the non-OECD, the assumption that the letters of credit have *less* than three months of remaining maturity results in another but similar identification scheme that is further described in Section 2.3.

[10] Khandelwal, Schott, and Wei (2013) follow a similar approach in a comparable empirical setting.

increase in the affected risk-weight category. This result implies a risk-weight elasticity of letter of credit-based share of exports of -0.0787 at the firm-country-product-level. For the non-OECD sample, the baseline regressions yield smaller yet comparable results: the decrease in risk-weights generates a 6.13% to 7.36% increase in export share for the affected rating categories. The related risk-weight elasticities -0.1226 and -0.1454 for letter of credit-based export shares. These results are robust to changes in empirical specifications through additional fixed effects or control variables, different clustering of the standard errors, or limiting the sample to specific product categories to better control for credit conversion factors. A placebo test with the fictitious Basel II adoption date of July 1, 2011 suggests that our results are not due to anticipation of new capital regulation's implementation.

However, these findings do not directly answer the question of whether Basel II adoption negatively affected trade flows as suggested in the financial press and banking surveys. To test for this possibility, we run export growth regressions. We find no discernable effect: even though firm-country-product level total exports aggregated over all payment types appear to respond negatively (positively) to the associated risk-weight increases (decreases), the observed effects are not statistically significant. Overall, our findings imply that the impact of changing risk-based capital requirements on international trade is subtler than the one suggested in the popular press or inferred from banking surveys.

Our work contributes to the emerging literature on the role of bank financing in international trade (e.g., Antras and Foley, 2015; Auboin and Engemann, 2012; Glady and Potin, 2011; Schmidt-Eisenlohr, 2013; Mateut, 2014; Niepmann and Schmidt-Eisenlohr, 2015). Our data and empirical approach allow us to focus on letter of credit export shares while controlling for firm-country-product unobservables. Findings here suggest that exogenous shocks to trade financing can affect firms' letter of credit-based share of exports. Our estimates are comparable to those found by Paravisini et al. (2014) who examine the *value* of Peruvian exports to financial shocks during the Great Recession. The results of this paper complement the findings of the recent strand of research on the impact of the Great Recession on international trade through the trade-finance channel (e.g., Ahn, Amiti, and Weinstein, 2011; Asmundson et al., 2011; Eaton, et al., 2011; Levchenko, Lewis, and Tesar, 2010; Chor and Manova, 2012; Paravisini et al., 2014) in two ways. First, we show that international trade reacts to changes in the implicit costs of processing letters of credit. This finding is consistent with both letter of credit pricing and rationing channels. Second, we find that, while letter of credit-based export shares were affected, the growth of total firm-country-product level exports were not affected by Basel II adoption. The latter finding suggests that concerns about Basel II adoption's impact on world trade is not supported in our data.

Our paper proceeds as follows. In Section 2 we provide information on (i) our data sources, (ii) the economic background, and (iii) our empirical specifications. Section 3 presents our empirical results (including robustness checks). Section 4 concludes the paper.

2. Data, economic background, and empirical specifications

2.1. Data

Our data come from two sources. Firm-country-product-trade financing level export flows data are obtained from the Turkish Statistical Institute (TSI).[11] Counterparty bank rating data are from BankScope. This section explains how we combine these data, and provides summary statistics that detail their heterogeneity.

To construct our dependent variable based on TSI data we proceed as follows. Given the Basel II adoption date of July 1, 2012, first we aggregate the monthly firm-country-product-trade financing-level export flows data available in the TSI dataset into annual data denoted by $t \in \{-1, 0, 1, 2\}$, each running from July 1 of year τ to June 30 of year $\tau+1$ with $\tau \in \{2009, 2010, 2011, 2012\}$. In a second step, we construct our dependent variable, firm-country-product letter of credit-based share of exports, by taking the ratio of letter of credit-financed flows within an annual period t with respect to all shipments by the same firm to the same country in the same HS6 product category (all export payment methods -- i.e., letters of credit, open account, and cash in advance -- combined) over the same period t. This dependent variable is first-differenced in order absorb confounding factors at the firm-country-product-level that are time-invariant during two consecutive annual periods. The resulting data is, then, split into two. Annual data based on periods $t=\{0, 1, 2\}$ are used in our main analysis: after we first-difference the dependent variable, we obtain a two-period panel with which we estimate our difference-in-difference regressions (see Section 2.3 for more detail).

Before proceeding further, we explain the reasons underlying the above-described choices. We aggregate monthly export flows into annual ones for the following reasons. First, we would like to attenuate the problems associated with zero-trade observations. Our dependent variable is the first-difference of letter of credit-based export share. This variable takes the value of zero when exporting firms use a type of financing other than letter of credit. Naturally, if there are no exports with any of the three trade-financing methods in a given year, we cannot calculate the letter of credit-based share of exports for that firm-country-product triplet in that period. Moreover, given that we first-difference the dependent variable, we require that there should be some exports in all periods for a firm-country-product triplet (irrespective of the trade financing method involved). Annual data are more amenable to such an empirical approach than observations with monthly or quarterly frequency. With the monthly or quarterly observations, the requirement to have at least some exports in all periods would lead to the loss of most of the data at our disposal. Second, we would like to account for the seasonality of the trade flows. Seasonality, if left unaccounted for, could affect the coefficient estimates of interest in our difference-in-differences model: we do not want the interactions of the ratings-range (i.e., risk-weights) and Basel II indicator variables to pick up what is in essence a seasonal variation in exports that arises

[11] These confidential data are accessed through dedicated computers at the TSI upon a special request and after a security clearance: statistics and regression results can only be exported upon approval of the TSI staff. We underline the fact that the TSI dataset has detailed export *flows* data but it has no export *shipments* data.

at the monthly or quarterly frequency. Seasonality could also induce serial correlation in the panel, for which one solution is indeed the aggregation of the data at the annual frequency (e.g., Bertrand, Duflo, Mullanaithan, 2004). Third, using share of letter of credit-based exports, rather than their values, as the dependent variable implicitly controls for unobservables that could affect firm-product-level exports, such as demand factors at the destination country.

We acknowledge that the TSI data, while very informative on export flows at the firm-country-product-payment term level, also have their weaknesses. First, the data do not contain any information on importers located in export-destination countries. So we cannot directly match exporters to their foreign importer-partners. That said, to the extent that the importer or group of importers in a given country to which a Turkish firm exports one particular product stays constant over two consecutive annual periods, any associated effects would be differenced-out in our empirical approach. Second, TSI data do not contain any information regarding letters of credit used in Turkish exports apart from the fact that the associated flows are financed with these instruments. As a result, we cannot observe the types or the maturities of the letters of credit involved for exports at the firm-country-product level. In our case, different types of letters of credit map into four different credit conversion factors. However, the credit conversion factor categories remain constant under both Basel I and Basel II (see Appendix Table A1). Nevertheless, we conduct additional tests for a product category for which the credit conversion factor is uniform. Another difficulty we face is that we cannot observe letter of credit maturities. To deal with this problem, we estimate our models using different risk-weight categories that depend on whether the said letters have more or less than three months of remaining maturity. Although there is no industry-specific evidence on the types of letter of credit used in international trade, findings of (domestic) trade credit research suggests that payment terms are industry specific (Ng, Smith and Smith, 1999). As a result, given the very detailed firm-destination-(HS6) product level data that we work with, we do not expect much variation over time in the types or maturities of letters of credit used in exports to a specific destination by a given firm. Moreover, we follow an empirical approach (see Section 2.3 below) that absorbs letter of credit type differences so long as they remain constant over two consecutive annual periods. Third, the TSI dataset does not include any information about the Turkish banks that hold the letters of credit at the request of their exporter-clients. That said, we can account for Turkish exporters' bank relations if the composition of the latter (i) remains constant over two consecutive annual periods (through first-differencing) or (ii) any changes therein would be captured by time-varying firm fixed-effects that we include. Finally, our data do not allow us to observe the identities of the counterparty banks (located in other countries) that have issued the letters of credit used in Turkish exports. As a result, we rely on a country-level proxy for the counterparty bank rating (please see next paragraph). These limitations, which are imposed on us by the nature of the TSI data, inevitably put some restrictions on the analyses that we can conduct (see the next two paragraphs). They also point to the importance of saturating our empirical models with many fixed-effects to absorb unobservables that otherwise might have confounding effects on our tests. In Section 3.3, we discuss

whether any of the dimensions for which we lack detail in the data at our disposal could somehow be driving the results that we obtain.

We combine the above-described TSI exports data with long-term agency ratings for banks. From the BankScope database we collect all available ratings by Fitch, Moody's, and S&P for individual financial institutions that could be issuing letters of credit for importer-clients of Turkish exporters.[12, 13] For many banks we have overlapping ratings from more than one of these three agencies, in which case we follow the rules imposed by the Turkish banking regulators:[14] if a foreign counterparty bank has two agency ratings, Turkish banks have to use the worst (lower) of the two ratings, if it has three ratings Turkish banks are required to use "the better of the worst two ratings" (i.e., the middle rating). Since the TSI data do not identify the letter of credit-issuing institutions, we use a proxy for the counterparty-bank rating at the export-destination country level. For this, we create a variable by weighting the observed bank ratings by the latest available total assets as well as the number of days the observed rating is valid for an institution (in case there are changes to that institution's ratings over the annual period) for a given country in a given annual period t. We weight the observed ratings for the following reasons. Some evidence suggests that a country's largest banks predominantly issue letters of credit: Del Prete and Frederico (2012) note that the issuance of trade guarantees (which include letters of credit) is limited to top-10 Italian banks which issue 78% of export- and 74% of import-related loans; whereas Niepmann and Schmidt-Eisenlohr (2015) observe that top-5 US banks account for 92% of all US trade claims (which, in their data, are mainly letters of credit, but also include other claims such as factoring and forfaiting). However, it is not clear to what extent Italy and the US are representative of all OECD countries, let alone non-OECD countries. For example, for Turkey (an OECD-member country), bankers that we spoke to indicated that while the largest Turkish banks typically issue letters of credit, smaller foreign banks are also locally active in this business (most likely thanks to the comparative advantage conveyed by them being owned by a larger foreign bank). As a result, our counterparty-bank rating proxy allows us to account for the fact that the largest banks in a country are most likely to be the most common issuers of letters of credit, without ruling out the possibility that small, specialized banks operating in the export-destination country might also be involved in this line of business.

The initial counterparty-bank agency rating coverage on BankScope have the following characteristics. For the OECD sample of 25 countries, the average (median) number of banks for which an agency rating can be observed is 85.28 (15), ranging from a minimum of 0 (Iceland, which does not

[12] Alongside Fitch, S&P and Moody's ratings, the Turkish regulator also allows ratings by JCR and DBRS agencies (e.g., BDDK FAQ number 68), but these agencies' ratings are not available in the Bankscope database.

[13] These include commercial banks, bank holding companies, state-, local government- or privately-owned savings banks, credit unions, cooperative banks, specialized government credit institutions (which include export-import banks), Islamic banks, and micro-finance institutions. We exclude the ratings of the following types of institutions: central banks, supranational entities (for ex., regional development banks), securities firms, investment banks, investment and trust corporations, and finance companies.

[14] As detailed in the BDDK Directive of June 28, 2012, Supplement 1, Section 2, Articles 1.5 and 1.6.

have any rated banks) to a maximum of 1,396 (Germany). The average (median) country-level total asset-weighted average counterparty-bank rating is approximately equivalent to Moody's A1 (A1), ranging from a maximum of Aa3 (New Zealand) to a minimum of Baa3 (Ireland). In the non-OECD sample, which consists of 133 countries, the average (median) country has 3.48 (0) banks rated, with a minimum of zero for 76 countries and a maximum of 98 (Russia). For the 57 non-OECD countries in our sample that have agency-rated financial institutions, the mean (median) country-level weighted-average counterparty bank rating is equivalent to Moody's Ba1 (Ba2) rating, ranging from a maximum of Aa3 (Malta and Singapore) to a minimum of B3 (Belarus and Ukraine).

We impose a number of restrictions on the combined TSI-BankScope data. First, we restrict ourselves to shipments by the manufacturing sectors, which formed 94% of Turkish goods exported in 2012. Second, we exclude firm-country-product level annual exports that are below 10,000 US $ to remove small shipments by marginal exporters that might otherwise influence our export share-based estimates. Third, given that our counterparty bank rating proxy is at the country-level, we require that the related firm-country-product-level observations remain within the same rating-range (e.g., within Aaa-Aa3 or A1-Baa3) that correspond to a particular risk-weight (see Table 1) throughout our sample. We do so by excluding countries for which the counterparty-bank rating proxy moves across rating ranges as defined by Basel II (e.g., from Aaa-Aa3 range into A1-Baa3, or vice versa), something which effectively fixes the credit ratings as of the pre-Basel II period. This restriction is needed in order to be able to estimate proper difference-in-differences models in which the rating-range (i.e., risk-weight) level confounding factors are appropriately captured by the same set of constants throughout the estimation period. Fourth, we exclude a number of countries because they are specific cases that might otherwise have an undue influence on our results: these include states that become OECD-members during our sample period, Cuba, so-called Arab Spring countries, together with Iran and United Arab Emirates.[15] After these restrictions we are left with exports by 9,085 firms to 25 OECD countries in 2,140 HS6 product categories and shipments by 11,419 companies selling 2,661 types of HS6-level

[15] Barbados, Belize, Grenada, Luxembourg, Pakistan, and Uruguay are excluded because the related weighted-average bank rating moves across rating-range (risk-weight) categories. Chile, Estonia, Israel, and Slovenia are excluded because these countries became OECD-members in 2010, during our sample period. Greece is excluded from the OECD sample because i) this is the only country with near-default ratings, ii) during the period studied the proxies for its banks' ratings (in terms of total assets-weighted average bank rating), although constant in mean (near default, with a 150% risk-weight category) change dramatically between categories. Hungary and Portugal are also excluded for a similar reason: i) they would be the only countries in the "junk" category with very few observations ii) although their yearly mean rating puts them in the speculative range (a 100% risk-weight category), the sovereign ratings worsen dramatically from the A1-Baa3 range according to Moody's to speculative grade. Cuba has no rated banks but is the only non-OECD country in sovereign default status throughout the sample period and a marginal export destination. Arab Spring countries' imports as well as their banks' credit ratings were negatively affected by social unrest and civil wars during our sample period. We drop Iran because it was subjected to an international embargo, which led to an increase in gold trade with Turkey to get around sanctions (Financial Times, February 18, 2013). We also eliminate United Arab Emirates because most of the unusual Turkish-Iranian gold transactions appear to have been done through this country (Financial Times, March 24, 2013).

products to 133 non-OECD countries. The data in our sample correspond to about 56% of Turkey's manufacturing exports during the July 2011-June 2013 period by value.

Before discussing the estimation results in the next section, first we go over some of the patterns in the data. Table 2 provides summary statistics for our dependent variable for the two annual periods (t=1 and t=2) around the adoption date of July 1, 2012. In Panel A, for the OECD sample, the average share for letter of credit-based exports is equal to 0.0363 pre-Basel II and to 0.0343 post-Basel II. The observed difference of 0.0030 is statistically different from zero at the 10%-level in a two-sided t-test. For the non-OECD sample, the average share for letter of credit-based exports is equal to 0.0584 before Basel II adoption and to 0.0550 after the application of the new rules. The difference of 0.0034 in the export shares is statistically different from zero at the 1%-level in a two-sided t-test. We underline the fact that these means are based on firm-country-product level share of exports. Our raw data (based on *value* of exports) exhibit patterns similar to those reported in the literature for other countries:[16] when aggregated over firms-products-destination countries the (value-based) fraction of Turkish trade that relies on letters of credit is approximately 5% for the OECD countries and roughly 15% for the non-OECD countries.

In fact, our export share data exhibit heterogeneity in many dimensions. One of these is the difference in the use of letters of credit across counterparty rating groups that form our "treatment" and control groups (based on the associated risk-weight categories). In Panel B of Table 2, we provide the average share for letter of credit-based exports for counterparty-bank rating-ranges that correspond to categories of risk-weights that are at the heart of our identification schemes. For the OECD sample, for countries whose banks are rated Aaa-Aa3 on average (and for which the corresponding risk-weight stays constant at 20%) the average share of exports is 0.0130 in the year prior Basel II adoption and 0.0118 in the year after (the difference of 0.0012 is not statistically significant in a t-test). Whereas for countries whose institutions are on average rated A1-Baa3 or non-rated (for which groups the risk-weight increases from 20% to 50%) the average share of exports decreases from 0.0455 pre-Basel II to 0.0433 post-Basel II (the difference of 0.0022 is marginally statistically significant at 11% in a two-sided t-test that allows for unequal variances). For the non-OECD sample, we see a larger share of exports based on letters of credit for higher-rated counterparties. For countries whose banks are rated Aaa-Aa3 on average (for which the risk-weights drop from 100% to 20%), Average share of exports is approximately 0.0958 pre-Basel II versus 0.0896 post Basel II (the difference is statistically insignificant). For countries with A1-Baa3 rated and non-rated institutions (for which risk-weight

[16] For example, Ahn (2014) states that letters of credit are involved in 5% of Columbian and 10% Chilean imports in 2011. Korean International Trade Association data indicate that 15% of South Korean exports involved letters of credit in 2012 (http://www.kita.net/statistic/index_eng.jsp). Niepmann and Schmidt-Eisenlohr (2015) report that 9.6% of U.S. goods exports (in terms of value) were settled this way. International Credits and Collections Survey conducted by the Finance, Credit and International Business Association (FCIB) indicate that in 2010, the exporter-reported country-level use of letter of credit as the main payment method has an average of 10.8% for a sample of 59 countries (Table 1 in Schmidt-Eisenlohr, 2013). Antras and Foley (2015) report that 6% of the exports of a large US frozen-poultry products producer are letter of credit-based.

decreases from 100% to 50%), the average share of letter of credit-based exports *decreases* from 0.0729 pre-Basel II to 0.0699 post-Basel II (the difference is statistically significant at the 10%-level in a two-tailed t-test). A decrease can also be observed for countries with Ba1-B3 rated counterparty banks (even though the related risk-weight remains constant at 100%) the average share of letter of credit-based exports is equal to 0.0337 pre-Basel II and 0.0299 post-Basel II (the observed difference of 0.0038 is statistically significant at the 1%-level in a two-sided t-test). While the observed decreases for the non-OECD sample appear counterintuitive (given that risk-weights also decrease), they are most likely due to confounding factors that are not controlled for in these summary statistics. For example, some of the observed patterns could be due to variations in the exports to different OECD countries by different industries, some of which might rely on letters of credit more than others (see, for example, Table 3).

To get a further understanding of the variation in our dependent variable, in Figure 1 we provide the frequency distributions for the share of exports based on letters of credit after excluding the zero-shares (to facilitate the reading of the graphs) in the annual period (t=1) that precedes Basel II adoption.[17] In Figure 1.a for the OECD sample, slightly more than 4% of the non-zero export share observations have an export share of 100%, suggesting that relatively few firms export a product to a destination country using letters of credit alone. In fact, for the OECD countries the frequency of lower use of letters of credit (related share of exports less than 10%) is more common (ranging roughly from 5% to 10% of the distribution) than higher use (share of export higher than 90%) of these instruments (ranging approximately from 2% to 4%). In Figure 1.b for the non-OECD sample, the frequency distribution of non-zero share of letter of credit-based exports is more evenly distributed (between 2% and 5%).

In Table 3 we provide statistics on the heterogeneity of export financing across industries, albeit at the level of groups of HS2-level industries (for the ease of exposition, given that we have 2,888 HS6-level products in the data). For example, in the year preceding Basel II adoption (i.e., during t=1), the letter of credit-financed exports account for only 0.42% of exports in the footwear and 0.65% in the "raw hides, skins, leather and furs" sectors, in contrast to 19.80% in the textiles, 17.41% in metals, and 16.48% in machinery industries.

To get a better sense of the heterogeneity observed in the data, in Table 4 we provide more information on the variation of the export patterns per firm, country, HS6-level product type, and combinations thereof, in the annual period prior to Basel II adoption (i.e., in t=1). In Table 4 panel A we examine the OECD sample. For exports based on letters of credit the average (median) of number of exporters per country is 66.3 (39), products per exporter 2.4 (1), products per country 72.8 (37), exporters per country-product pair 1.9 (1), products per exporter-country pair 2.0 (1), and export-destination countries per firm-product pair 1.4 (1). For exports to OECD countries that are based on

[17] This exclusion is done to facilitate the reading of the frequency distributions: in the annual period t=1 immediately preceding Basel II adoption, the percentage of observations for which the letter of credit share of exports is equal to zero is 92.8% for the OECD sample and 91.1% for the non-OECD sample.

other types of financing (i.e., open account or cash in advance), the average (median) of number of exporters per country is 935 (645), products per exporter 2.5 (1), products per country 526.6 (445), exporters per country-product pair 3.33 (1), products per exporter-country pair 1.9 (1), and export-destination countries per firm-product pair 2.0 (1).[18] We observe similar differences along these dimensions for the non-OECD sample. The statistics presented in Table 4's panels A and B underline the importance of controlling for heterogeneity inherent in the exports data.

The above observed differences in the use of letters of credit for the financing of international trade is likely to depend on many firm characteristics, some of which may be directly observable (e.g., industry segment as in Hoefele et al., 2016), others adequately proxied by measurable firm characteristics (e.g., firm size) or simply unobservable (e.g., the bargaining power of the firm in negotiating with its importers as in Demir and Javorcik, 2014). Some of the observed heterogeneity might reflect unobservable qualities of the exported good (for example, the same manufacturer can ship different quality versions of the same product to two different destination countries as in Manova and Zhang, 2011). Given the limitations of the data at our disposal, we do not resort to Heckman-type selection models to adjust for the presence of unobservable characteristics at various levels, but instead we rely on first-differences to absorb confounding factors at the firm-country-product level.[19]

Next, we focus on the time-trends in our dependent variable across counterparty rating groups that correspond to risk-weights that are the focus of our identification strategy. We would like to test the effects of changes in the risk-weights (hence the cost of capital) for letters of credit on the shares of letter of credit-based exports using a difference-in-differences approach based on the interactions between rating-range and Basel II indicator variables. Such an approach requires that the basic difference-and-differences assumption hold in our data: i.e., the pre-treatment paths of the share of letter of credit-based exports should be similar across rating-range group for which the risk-weights change with Basel II compared to that for which they remain the same. To examine whether this is the case, first we plot the average share of letter of credit based exports (at the firm-country-product level) for counterparty rating-ranges that correspond to risk-weights listed in Table 1. Figure 2 shows that the basic difference-in-differences assumptions are unlikely to prevail in either of the two samples (OECD and non-OECD) that we use. There are pre-Basel II differences not only in the levels of share of exports (which would be absorbed by the related rating-range indicator variables), but more importantly, in the trends of share of letter of credit-based exports between the treated groups (for which the risk-weights change) and control groups (for which the risk-weights remain the same).[20]

[18] In this case, given the level of detail of our data, we combine all non-letter of credit financed exports, including data from firms that rely on letters of credits for their other exports besides the firm-product-country combination at hand.

[19] Put differently, any firm-level selection correction that would be introduced in our export share regressions would be differenced-out in our baseline regressions at the firm-country-product level.

[20] The observed differences in time-trends may arise, at least in part, due to the fact that Turkish banks were asked to conduct Basel II Quantitative Impact Studies (QIS) by the Turkish regulators, as suggested by the Basel Committee. These QIS would have drawn the banks' attention to various costs associated with Basel II adoption:

These observations suggest that our empirical approach has to take into account the differences in the time trends across the treatment and control groups for the letter of credit-based share of exports. As a result, we cannot implement a difference-in-differences approach in the *level* of share of exports, since the treated and the control groups do not appear to react the same way, on average, in the absence of the treatment in the pre-Basel II period. To deal with this problem, as we explain in more detail in Section 2.3 below, we follow Khandelwal, Schott, and Wei (2013) and implement difference-in-differences with *first differenced* (rather than the *level* of) share of exports.[21] But before explaining our empirical approach, first we discuss why Turkey is an interesting case to study.

2.2. Economic background

The way in which Basel II was implemented in Turkey, when combined with the availability of detailed exports data, provides us with a particularly suitable quasi-natural experiment to examine the impact of risk-based capital requirements on international trade. Turkish banking regulators required that Basel II be implemented as of July 1, 2012 *only* in its Standardized Approach form.[22] , which rules out the Internal Rating Based (IRB) Approach typically adopted by large banks. One consequence is that we do not have to worry about the identification-related complications that the latter approach (in its Foundation or Advanced version) brings: banks' internal counterparty risk assessments, which are proprietary and hence typically not accessible to researchers, tend to differ across institutions in the capital charges that they imply for a given on- or off-balance sheet position (see, for example, Financial Times, February 26, 2013).[23] In contrast, the components of the Standardized Approach of Basel II are public knowledge: the same set of identification schemes applies to all banks operating in Turkey as of July 1, 2012.

Another reason to consider Turkey is that the country is economically relevant for international trade. It is a member of the OECD, WTO and the G-20. As of 2012, the country was the world's 17th largest economy, 22nd largest exporter by value (15th largest exporter in manufactured goods that we examine), and 14th largest importer.[24] Turkey is in a customs union for manufactured goods with the EU since 1996. It is the fifth largest exporter to this economic zone (sixth largest in manufactured goods)

banks might have started to implement some, if not most, of the Basel II-triggered policy changes earlier than the adoption date of July 1, 2012, potentially generating changes in time trends.

[21] Consistent with the Khandelwal, Schott, and Wei (2013) approach, we observe no difference between the treatment groups (for which the risk-weights change) and control groups (for which they do not) when we re-run the cross-sectional regressions with double-differenced data (see even-numbered columns of Appendix Table A3).

[22] Starting with January 1, 2014 Turkey adopted Basel III, which is being implemented in stages until 2020.

[23] Turkish banking authorities made it clear that the internal rating-based approach would eventually be introduced and asked the banks in their jurisdiction to develop their own internal rating models. But, as of July 2013 no Turkish bank was permitted to use the IRB approach officially.

[24] These rankings treat the E.U. as a single economy consisting of 27 member-country economies. Ranking based on the size of the economy according to the 2012 estimates of the International Monetary Fund (IMF). Rankings based on trade according to the 2011 estimates of the World Trade Organization (WTO).

and its seventh largest importer.[25] Moreover, the manufactured goods that we examine formed approximately 94% of total Turkish exports of goods in 2012. So our inferences are based on a large, diversified economy that is relevant for global trade, exporting overwhelmingly manufactured goods.

However, the economic background of the Turkish case and the available data also present certain challenges. First, Turkish banks, unlike their EU or US counterparts during the same period, were well capitalized as of June 2012. Turkish banks' risk-weighted Tier 1 plus Tier 2 capital ratio was more than twice the amount required by Basel II: 16.47% as of July 1, 2012.[26] Although Basel II led to a decrease in Turkish financial institutions' risk-based capital ratios, the average effect was approximately a 1.5% drop, leaving the banking sector capitalized at roughly 15%. If Turkish banks chose to internalize the capital charges resulting from Basel II adoption (to their benefit in those cases when capital charges decreased and to their clients' benefit when the implicit costs increased) we would be less likely to detect changes in the related trade shares.

Second, Turkish banks need not fully reflect Basel II related changes to capital charges into their prices for letter of credit-clearing. Although some of the bankers that we spoke to indicated that the prices of letter of credit-related services were affected by Basel II adoption, others stated that large-corporate customers doing repeated business with their institution were less likely to be affected compared to small and medium sized enterprises (SMEs) with less frequent export transactions. Moreover, the overall effect around Basel II adoption would depend on Turkish banks' net exposure to letters of credit in different rating ranges given the domiciliation (OECD or non-OECD) of the counterparty banks. This is because some of the risk-weight changes (say, for A1 to Baa3 or non-rated OECD-domiciled bank counterparties) would require holding more capital, but others (say, for A1 to Baa3 or non-rated non-OECD-domiciled bank counterparties) would need less capital.

Finally, there may be wider Basel II related effects in the economy that could have affected Turkish firms' exports indirectly. Higher risk-based capital requirements may lead banks to ration credit or adjust its terms differentially across groups of firms (for example, large firms versus SMEs). Such changes might affect production in general through lower availability of capital for investment and working capital. Because exports typically rely on working capital over longer periods (due to longer shipment periods involved in international trade), such general Basel II effects could have confounding effects that need to be accounted for in our empirical set-up. In Section 4.3 we discuss in detail whether one or more of the aforementioned dimensions mentioned in this sections could be driving our empirical results. Next, we explain the empirical models that we use.

[25] The E.U.-Turkish customs union does not cover agriculture or the services sector. This said, manufactured goods include processed food items.

[26] As another (but partial) indication of capital strength, we should note that after the adoption of higher capital requirements under Basel II, the S&P did not change its ratings for five large Turkish banks that it evaluates.

2.3. Empirical specifications

To conduct our analysis, we estimate a difference-in-differences model in which the dependent variable is the first-differenced letter of credit-based share of exports at the firm-country-product level. We calculate the letter of credit-based export shares as $S_{f,c,p,t} = EXPORTS\_LC_{f,c,p,t} / EXPORTS\_TOTAL_{f,c,p,t}$ where, $S_{f,c,p,t}$ is the share of letter of credit-based exports for firm *f* exporting (HS6-level) product *p* to destination country *c* in period *t* with respect to total exports of firm *f* in product *p* to country *c* during time *t*; *EXPORTS* denotes value of exports, suffixes *_LC* and *_TOTAL* denote exports based on letters of credit and total exports, respectively. Using the first-differenced share of letter of credit based exports allows us to account for unobservables at the firm, destination country, product levels, or any combination thereof, without having to resort to large numbers of fixed effects that would be needed if we were to use the level of export shares instead: the first-difference of the dependent variable absorbs any firm-country-product level confounding factors that remain constant over two consecutive annual periods. For example, for the OECD (non-OECD) sample any product-demand effects would be absorbed for each of the 2,140 (2,661) manufactured goods that can be shipped to 25 (133) different countries by any one of 9,085 (11,419) Turkish exporters. This level of detail allows us to control not only for general country-level demand for Turkish goods or overall demand for a particular product of a given firm, but also for subtler effects, such as differences in demand due to product quality that varies by destination countries for a given exporter (e.g., Manova and Zhang, 2011).

Moreover, as discussed in the end of Section 2.1 above, the standard difference-in-differences assumptions regarding the trend of the dependent variable being the same (on average) for both the treated and control groups do not hold for the OECD and non-OECD samples (see also section 4.3 below). To properly deal with this problem we follow Khandelwal, Schott and Wei (2013) and use three periods of annual data denoted by the subscript t, which, when first-differenced, allow us to run a difference-in-differences model with a two-period panel of differenced observations for the same firm-country-product triplet. The resulting specifications, which are also called the "triple-differences" models, capture the same treatment effect as a difference-in-differences model, but allow for differences in pre-treatment trends for the treated and the control groups (e.g., Mora and Reggio, 2013; and Lee, 2016).

First, let's consider difference-in-differences regression for the OECD sample, which corresponds to the identification scheme that involves a move from column 1 to column 4 of Table 1:

$$\Delta S_{f,c,p,t} = \beta_1 \, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t} \qquad (1)$$

where Δ denotes the first-difference operator (such that $\Delta S_{f,c,p,t} = S_{f,c,p,t} - S_{f,c,p,t-1}$); subscript *t* denotes annual periods with $t \in \{1, 2\}$; $D\_A1\text{-}Baa3\&NR_c$ is equal to one if the letter of credit-issuer counterparty-banks in the destination OECD country *c* have throughout our sample a long-term credit rating between A1 to Baa3 according to Moody's (A+ to BBB- according to S&P or Fitch) or none of them are rated,

and for which the risk-weight increases from 20% to 50% with Basel II, and zero otherwise; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise; $\gamma_{c,p}$ and $\gamma_{p,t}$ denote country×product and product×time fixed-effects, respectively; and ε is the error term of the OLS regression.[27] The omitted (i.e., the base case) category comprises exports involving letters of credit issued by counterparty banks with average ratings between Aaa to Aa3 according to Moody's (AAA through AA- according to S&P or Fitch) for which the risk-weight for OECD-country domiciled banks is 20% under *both* Basel I and II. The empirical model does not include rating categories below Baa3 because there is no OECD country whose banks' average rating is equal to Ba1 or below consistently throughout the sample period. Actually, the fact that there are no countries whose banks were consistently rated Ba1-B3 or Caa1 and below effectively constrains OECD regressions to the case for which the remaining maturities of letters of credit are longer than three months (i.e., to a move from column 1 to column 4 in Table 1).

The coefficient estimate of interest is β_1, which captures the effect of Basel II on export shares due to Basel II "treatment" after which the risk-weight for letters of credit issued by the OECD-based A1 through Baa3-rated or non-rated counterparty banks increased from 20% to 50% (compared to the "non-treated" letters of credit issued by OECD-country counterparty banks for which it remained constant at 20%). Our hypothesis suggests that β_1 should be negative: as the capital charge for letters of credit became 150% more expensive, the associated exports, whose financing involves letters by A1 to Baa3 rated counterparty banks located in the OECD, are expected to decrease in the post-Basel II period.

We estimate a model similar to Eq. (1) for the non-OECD sample. As we cannot observe letter of credit maturities, we first assume that the average letter of credit has a remaining maturity of *longer* than three months (which corresponds to a move in Table 1 from column 2 to 4):

$$\Delta S_{f,c,p,t} = \beta_1\, D\_Aaa\text{-}Aa3_c \times D\_BASELII_t + \beta_2\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t} \quad (2)$$

where, $D\_Aaa\text{-}Aa3_c$ is equal to one if the letter of credit-issuing counterparty-banks in the destination non-OECD country *c* have, on average, ratings that are between Aaa to Aa3 (for which risk-weight drops from 100% under Basel I to 20% under Basel II) throughout the sample period, and zero otherwise; $D\_A1\text{-}Baa3\&NR_c$ is equal to one if destination counterparty banks have instead a rating between A1 and Baa3 on average in non-OECD country *c* or they are not rated by any of the three rating agencies between July 2011 and June 2013 (for which groups the risk-weight drops from 100% to 50%), and zero otherwise; with the remaining variables being as described above. The base-case (omitted) category involves countries whose banks are rated Ba1 through B3 on average. For the non-OECD sample, the expected signs of the coefficient estimate of interest β_1 (and β_2) are now both positive: as the risk-weight applied to a letter of credit from a counterparty bank rated in the Aaa to Aa3 (A1 to Baa3 or non-rated) range decreases from 100% to 20% (50%), our hypothesis suggests that related

[27] In equations (1) through (3), coefficient estimates for the stand-alone indicator variables $D\_A1\text{-}Baa3_c$ and $D\_BASELII_t$ are absorbed into $\gamma_{c,p}$ and $\gamma_{p,t}$, respectively.

export shares would increase with respect to the base-case category (Ba1 to B3 rated counterparties) for which risk-weight remains constant at 100%. The channels behind the expected sign of β_1 and β_2 remain the same: a positive coefficient estimate would be consistent with a pricing channel (clearing of letters becoming cheaper as risk-weights go down) and/or a "rationing" channel (Turkish banks would be more open to increasing their exposures to counterparty banks for which risk-weights decrease).

To account for the possibility that export-letters of credit can have, on average, remaining maturities *less* than three months, we also estimate a version of Eq. (2) for non-OECD countries after a slight modification (this case corresponds to a move from column 2 to column 3 in Table 1):

$$\Delta S_{f,c,p,t} = \beta_1 \, D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t} \tag{3}$$

where, $D\_Aaa\text{-}Baa3\&NR_c$ is equal to one if the letter of credit-issuing counterparty-banks in the destination non-OECD country c have, on average, ratings that are investment-grade or non-rated (for both of which risk-weight drops from 100% under Basel I to 20% under Basel II) throughout the sample period, and zero otherwise; with the remaining variables being as described above. For this non-OECD sample, the expected sign of the coefficient estimate of interest β_1 is still *positive*: as the risk-weight applied to a letter of credit from a counterparty bank rated investment grade decreases from 100% to 20%, our hypothesis suggests that related export shares would increase with respect to the base-case category (Ba1 to B3 rated counterparties) for which risk-weight decreases from 100% down to 50%.

We also define a regression equation for the pooled OECD and non-OECD samples (after assuming the letters of credit have a remaining maturity of *longer* than three months):[28]

$$\begin{aligned}\Delta S_{f,c,p,t} = {} & \beta_1 \, D\_RW\_INCREASE_c \times D\_BASELII_t + \beta_2 \, D\_RW\_DECREASE_c \times D\_BASELII_t \\ & + \beta_3 \, D\_OECD_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}\end{aligned} \tag{4}$$

where $D\_RW\_INCREASE_c$ is an indicator variable that is equal to one if the risk-weight increases with Basel II adoption (irrespective of OECD or non-OECD destination countries), and zero otherwise; $D\_RW\_DECREASE_c$ is an indicator variable that is equal to one if the risk-weight decreases (irrespective of OECD membership of the export-destination country), and zero otherwise; D_OECD_c is an indicator variable that is equal to one if the export-destination is an OECD-member country; and $D\_BASELII_t$ being defined as above. As such, $D\_RW\_INCREASE_c$ covers exports to OECD-member countries for which the letter of credit risk-weight increases from 20% to 50% (i.e., OECD countries whose average counterparty-bank rating is between A1-Baa3). In contrast, $D\_RW\_DECREASE_c$ only captures exports to non-OECD countries for which the letter of credit risk-weight decreases from 100% to either 20% or 50% (i.e., non-OECD countries whose average counterparty bank rating range is

[28] We would like to thank the editor for suggesting this formulation.

covered under $D\_Aaa\text{-}Baa3\&NR_c$ defined above). $D\_OECD_c \times D\_BASELII_t$ is needed to capture the post-Basel II differences in the base-cases.[29]

While the above specifications (1) through (4) allow us to detect any changes to letter of credit based export *shares* (at the firm-country-product level) following Basel II adoption, they cannot help us to detect whether value of overall exports were affected at the same level of detail. This is an important point: if aggregate exports (by value) were to be affected firm-country-product level by Basel II adoption, then the trade organizations' concerns, which are indicated in the Introduction, would be warranted. An alternative scenario would be one under which firm-country-product level aggregate exports would not be affected because exporters would seek and obtain changes to export settlements by changing payment types (i.e., moving across letter of credit and non-letter of credit terms). In any case, the effect, if any, should be more pronounced for firms with a higher usage of letters of credit. We test for this possibility by running the following regression for the OECD sample:[30]

$$\begin{aligned}\Delta ln(EXPORTS\_TOTAL_{f,c,p,t}) = {} & \beta_1\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t \times S_{f,c,p,t=1} \\ & + \beta_2\, D\_A1\text{-}Baa3\&NR_c \times S_{f,c,p,t=1} \\ & + \beta_3\, D\_BASELII_t \times S_{f,c,p,t=1} + \beta_4\, S_{f,c,p,t=1} + \gamma_{c,p,t} + \varepsilon_{f,c,p,t} \qquad (5)\end{aligned}$$

where, $ln(EXPORTS\_TOTAL_{f,c,p,t})$ is the natural logarithm of the value of total exports (with all types of export financing combined) of firm f to country c for a given product p; $S_{f,c,p,t=1}$ is the pre-Basel II (i.e., period t=1) share of letter of credit financed exports for firm f to country c for a given product p; with the other variables being defined as above. For the non-OECD sample, we define a similar growth regression equation:

$$\begin{aligned}\Delta ln(EXPORTS\_TOTAL_{f,c,p,t}) = {} & \beta_1\, D\_Aaa\text{-}Aa3_c \times D\_BASELII_t \times S_{f,c,p,t=1} \\ & + \beta_2\, D\_A1\text{-}Baa3_c \times D\_BASELII_t \times S_{f,c,p,t=1} \\ & + \beta_3\, D\_Aaa\text{-}Aa3_c \times S_{f,c,p,t=1} + \beta_4\, D\_A1\text{-}Baa3_c \times S_{f,c,p,t=1} \\ & + \beta_5\, D\_BASELII_t \times S_{f,c,p,t=1} + \beta_6\, S_{f,c,p,t=1} + \gamma_{c,p,t} + \varepsilon_{f,c,p,t} \qquad (6)\end{aligned}$$

In the next section, we present our main findings.

3. Main results

Before discussing the results, we note that the regression equations, which are estimated with Ordinary Least Squares (OLS), involve country×time clustered standard errors. This is because our key variation of interest (the interaction of the indicator variable that traces the country-level average credit

[29] As in equations (1) through (3), country×product fixed-effects would soak-up the stand-alone D_OECD_c indicator variable, whereas product×time fixed-effects would fully account for the stand-alone $D\_BASELII_t$.

[30] We thank an anonymous referee for suggesting these specifications and tests on total export volumes.

quality of the counterparty bank with the Basel II indicator variable) is at the country-time level (e.g., Bertrand, Duflo, Mullanaithan, 2004). We have 25 OECD and 133 non-OECD export-destination countries and two periods (after first-differencing the data). In this setting, the number of truly independent clusters of data in each period of our two-period panel is 50 for the OECD and 266 for the non-OECD sample: firm-product pairs in the same country-year will face correlated shocks from common business cycles that are country specific. As part of our robustness checks, we also estimate our baseline regressions under alternative clustering options (see Section 4.3).

3.1. Estimates for the sample of OECD countries

First, we focus on the OECD sample, which is limited to exports to countries that are members of this organization throughout the sample period and whose banks have weighted-average ratings that are either investment grade or non-rated given the imposed data restrictions.[31] This means that we can only conduct tests under the assumption that the letters of credit have a remaining maturity more than three months, as OECD countries whose banks have on average below investment grade ratings drop out of our sample due to the implemented data constraints.

The findings are presented in Table 5. The result for Eq.(1), our baseline regression, are presented in column 1 of Table 5. The estimate of the coefficient of interest β_1 for the interaction $D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t$ is equal to -0.00537 (i.e., -0.537 percentage points), which is statistically significant at the 1%-level. This is an economically significant effect, given that in Table 2, the pre-Basel II share of letter of credit-based exports related with OECD counterparty banks for which the risk-weight increases is equal to 0.0455 (i.e., 4.55 percentage point) on average: the share of exports with letters of credit issued by A1-Baa3 rated counterparty banks decrease by -11.80% (=-0.00537/0.0455) as risk-weights increase with Basel II. This suggests that the risk-weight elasticity of letter of credit-based share of exports is equal to -0.0787 (=[-0.00537/0.0455]/[(0.50-0.20)/0.20]) at the firm-country-product-level as the risk-weights increase 150% from 0.20 to 0.50. In other words, for the OECD sample, a 10% increase in risk-weight leads to an approximately 0.8% decrease in letter of credit-based export-share for firms, after controlling for firm-country-product level confounding factors.

First-difference of the share of letter of credit-based exports takes out any confounding effects that are at the firm-country-product level so long as such effects remain constant between annual periods t=0 and 1 or t=1 and 2, separately. Moreover, we saturate the baseline model with country×product and product×time fixed-effects (all at the HS6-level) that would account for other potential confounding factors (such as country-level demand for a particular Turkish export product, or aggregate demand for a particular product over time). Nevertheless, this approach may not remove general Basel II effects that could affect some or all exporters in t=2. These could arise, for example, because higher risk-based capital standards may lead to less lending to all or some borrowers (e.g., small and medium sized

[31] The only OECD-member country whose banks are not rated is Iceland.

enterprises), because exports have longer payment-cycles and hence require more working capital. Less bank lending after Basel II, when combined with risk-weight changes, might affect the composition of exports depending on their types of trade-financing. It could also be that there are more changes to exporter-bank relations in Turkey around Basel II. To account for these and other plausible indirect Basel II effects, we re-estimate Eq. (1) after adding using time-varying firm fixed-effects to our baseline regression: in column 2 of Table 5, after the addition of firm×time fixed-effects, the coefficient estimate for *D_A1-Baa3×D_BASELII* is equal to -0.00588, which is statistically significant at the 1%-level. In percentage terms, as the risk-weights increase from 0.20 to 0.50 (a 150% change) for the A1-Baa3 rated counterparty banks, the letter of credit-based export share decreases by 12.92% (= -0.00588/0.0455). The associated elasticity of share of letter of credit-based exports to changes in risk-weights is equal to -0.0862 (=-0.1295/1.50): a 10% increase in the risk-weight leads to roughly 0.9% drop in the associated share of exports at the firm-country-product level. These results are very similar to the ones based on the β_1 estimate of column 1 of the same table. They show that the change in the share of exports using letters of credit is a *within-firm* phenomenon. We conclude that for the OECD sample, any indirect Basel II effects do not materially influence our results.

Next, we estimate Eq.(1) with better controls for time-varying firm-product confounding factors (such as time-varying demand for individual firms' products). To do so we replace the fixed-effects (combination of the product×time and firm×time) of column 2 with the firm×product×time fixed-effects, with the products at the HS2-level.[32] The coefficient estimate for the interaction *D_A1-Baa3×D_BASELII* is not materially affected by the addition of this triple-interaction fixed-effect: in column 3 of Table 5, the estimate for β_1 is equal to -0.00592, statistically significant at the 1%-level, a result that is similar in magnitude to those of columns 1 and 2 of the same table.

3.2. Estimates for the sample of non-OECD countries

In Panel A of Table 6 we present estimates of Eq. (2), which assumes that the letters of credit have, on average, maturities *longer* than three months. This assumption leaves us with two sets of counterparty-bank rating ranges: Aaa through Aa3 for which the risk-weight decreases from 100% to 20%, and A1 through Baa3 for which the risk-weight decreases from 100% to 50%. In column 1, the coefficient estimate for the interaction *D_Aaa-Aa3×D_BASELII* is equal to 0.0014 but statistically insignificant. In fact, the coefficient estimates for *D_Aaa-Aa3×D_BASELII* are never statistically significant in columns 2 through 3 of Table 6 Panel A. These results might be due to the fact that the group of countries whose banks are, on average, rated Aaa through Aa3 include three marginal export-

[32]In this specification country×product fixed effects are still defined at the HS6-level. For firm×product×time fixed-effects we limit ourselves to the HS2-level product categories because at the HS6-level the total number of fixed-effects (i.e., country×product and firm×product×time all combined together) would have amounted to 63% of OECD observations and 73% of the non-OECD observations.

destination countries (Hong Kong, Malta and Singapore).[33] Next, we focus on the group of non-OECD countries whose banks are on average rated A1 through Baa3 or not rated in our sample. For this rating range the risk-weight decreases from 100% to 50% (a 50% drop) with Basel II adoption. In column 1 of Table 6 Panel A the coefficient estimate for the interaction *D_A1-Baa3×D_BASELII* is equal to 0.00447, which is statistically significant at the 5%-level. This result points to a 6.13% increase in export share for the letters of credit from counterparties rated A1 through Baa3, given that in Table 1 the average fraction of letter of credit-based exports is equal to 0.0729 for the same group. The associated risk-weight elasticity of letter of credit-based export share is equal to -0.1226 (= [0.00490/0.0729] / [(0.50 - 1.00)/1.00]): a 10% decrease in the cost of letter of credit (due to a 10% decrease in the risk-weight) leads to a 1.23% increase in the letter of credit-based export share after having controlled for firm-country-product level confounding effects. This elasticity of -0.1226 for the non-OECD sample is larger than the one we obtained for the OECD sample (-0.0787) for the same rating range (A1 through Baa3), suggesting that, holding counterparty rating-range constant, share of letter of credit based exports for the non-OECD are more responsive to changes in risk-weights. It could also be that there is asymmetry between cost increases and cost decreases, a possibility that we test for in Section 3.3.

In column 2 of Table 6 Panel A, we include firm-time fixed-effects into Eq.(2) and we obtain results that are similar to those of the whole sample. The coefficient estimate for *D_A1-Baa3×D_BASELII* is equal to 0.00530, which is statistically significant at the 5%-level. Given that the average share of letter of credit-based exports to countries with A1-Baa3 rated banks is equal to 0.0729, the coefficient estimate of 0.00530 implies an increase of 7.27% (=0.00551/0.0729). The corresponding risk-weight elasticity of letter of credit-based export share is equal to -0.1454 (= [0.00551/0.0729] / [(0.50 – 1.00)/1.00]). This elasticity suggests that, when the risk-weight (cost) associated with the letter of credit decreases by 10%, the share of exports to countries whose banks are rated A1-Baa3 or are non-rated increases by 1.45% compared to share of exports by the same group of firms to countries whose banks are rated Ba1-B3. Another interpretation of this result is that the effects that we observe are not due to general Basel II effects (such as changes in lending standards), which in this case would be soaked up by the time-varying firm fixed effects. Yet another interpretation would be that firm-level time-varying unobservables (such as the changes in an exporter's bank relations) do not appear to have a significant bearing on our results. In column 3 of Table 6 Panel A, we replace *product×time* fixed-effects of Eq.(2) with the stronger *firm×(HS2-level) product×time* fixed-effects, but our estimates are barely affected: *D_A1-Baa3&NR×D_BASELII* coefficient estimate with these new set of fixed-effects is equal to 0.00519, which is statistically significant at the 10%-level.

In Table 6 Panel B, we examine whether the results for the non-OECD sample still hold under the assumption that the export letters of credit have, on average, remaining maturities *shorter* than three

[33] In 2012, Malta received 0.60% of Turkey's total exports (by value), Singapore 0.29%, and Hong Kong 0.22%.

months by estimating Eq.(3). In column 1, the coefficient estimate for the interaction *D_Aaa-Baa3&NR × D_BASELII* is equal to 0.00439, which is statistically significant at the 5%-level. The observed impact of risk-weight change is not trivial economically: it amounts to an 5.96% increase in export share, which is equal to 7.36% for investment grade plus non-rated non-OECD countries. The corresponding risk-weight elasticity of letter of credit-based export share equals -0.0746 (= [0.00439/0.0736] / [(0.20 - 1.00)/1.00]).

In column 2 of Table 6's Panel B, we re-estimate Eq.(3) after having introduced firm-time fixed-effects to soak up any firm-level confounding effects: the coefficient estimate for *D_Aaa-Baa3&NR×D_BASELII* is equal to 0.00497, which is statistically significant at the 5%-level. In column 3 of Table 6 Panel B, we re-estimate Eq. (3) after introducing *firm×(HS2-level) product×time* fixed-effects instead of *product×time* fixed-effects: the coefficient estimate for the interaction *D_Aaa-Baa3&NR × D_BASELII* is 0.0482, which remains statistically significant at the 10%-level. In the next section, we present the estimates of Eq. (4) using a pooled OECD and non-OECD sample.

3.3. Estimates for the pooled (OECD plus non-OECD) sample

Pooling together the OECD and non-OECD samples has a number of advantages. First, combining the two samples allows us to preserve degrees of freedom, leading to more efficient estimation given the very large number of fixed-effects in our empirical models. Second, the *product×time*, *firm×time*, and *firm×(HS2-level)* p*roduct×time* fixed-effects would be more efficiently estimated (as they would be estimated with more data available in the pooled sample). The estimates of Eq. (4) are presented in Table 7, under the assumption that the letters of credit have longer than three months of maturity on average.[34] In column 1, the coefficient estimate for $D\_RW\_INCREASE_c \times D\_BASELII_t$, which corresponds to the A1-Baa3 and non-rated counterparties located in OECD countries (as traced by our country-level proxy) that are subjected to an increase in risk-weights (from 20% to 50%), is equal to -0.00512, which is statistically significant at the 1%-level. This estimate is very similar to the coefficient estimate of -0.00537 in column 1 of Table 5. Still in column 1 of Table 7, the coefficient estimate for $D\_RW\_DECREASE_c \times D\_BASELII_t$, which corresponds to the combined cases of Aaa-Aa3, A1-Baa3 and non-rated non-OECD-based counterparty banks, is equal 0.00457 and statistically significant at the 5%-level. Again, this estimate is very similar to the estimate of 0.00447 for the *D_A1-Baa3&NR×D_BASELII* interaction in column 1 of Table 6.

The pooled sample regression gives us the possibility to test for the equality the coefficient estimates of $D\_RW\_INCREASE_c \times D\_BASELII_t$ and $D\_RW\_DECREASE_c \times D\_BASELII_t$, which are provided at the bottom of column 1 in Table 7. We reject the null hypothesis that the coefficient estimate

[34] Estimating Eq. (4) under the assumption that the letters of credit have *shorter* than three months would amount to estimating only the $D\_RW\_DECREASE_c \times D\_BASELII_t$ interaction but with a larger sample: $D\_RW\_INCREASE_c \times D\_BASELII_t$ cannot be estimated since there is no corresponding change in the risk-weights in this case (see Table 1).

of -0.00512 for $D\_RW\_INCREASE_c \times D\_BASELII_t$ equals the coefficient estimate of 0.00457 for $D\_RW\_DECREASE_c \times D\_BASELII_t$: the Wald test statistic is equal to 18.89, which statistically significant at the 1%-level. In a second step we test whether the observed effect is symmetric, i.e., we test for the equality of the absolute values of the observed coefficient estimates: the test statistic is equal to 0.0678, which is not statistically significant. So the observed effect Basel II adoption on letter of credit based exports appears to be of similar magnitudes even if of different directions for the risk-weight increases versus decreases.

We make similar general observations in column 2 (with the addition of *firm×time* fixed-effects to Eq. (4)) and column 3 (with the addition of and *firm×(HS2-level)* p*roduct×time* fixed-effects to Eq. (4)). For example, in column 3 of Table 7, with the stronger *firm×(HS2-level)* p*roduct×time* fixed-effects, the coefficient estimate for $D\_RW\_INCREASE_c \times D\_BASELII_t$ is equal to -0.00448 (which is statistically significant at the 1%-level) and the one for $D\_RW\_DECREASE_c \times D\_BASELII_t$ is equal to 0.00539 (which is statistically significant at the 5%-level). And again, as reported at the bottom of column 3 of Table 7, we reject the null hypothesis for the equality of these observed coefficient estimates, but we cannot reject the null hypothesis of the equality of their absolute values.

While pooled sample estimates are line with our estimates using separate OECD and non-OECD samples, these results do not clarify the actual impact, if any, of Basel II associated risk-weight changes on total trade, something which we examine next.

3.4. Trade growth regressions

In Table 8 we present the estimates of trade growth regressions, for the OECD and non-OECD samples, respectively, in which we test whether Basel II actually affected total trade (aggregated over all payment terms) at the firm-country-product level. In column 1 of Table 8, Eq. (5) for the OECD total trade growth, our test relies on the triple interaction $D\_A1\text{-}Baa3_c \times D\_BASELII_t \times S_{f,c,p,t=1}$ for which the coefficient estimate is equal to -0.0557 but not statistically significant. Despite having the correct negative sign, our test indicates that, for higher letter of credit export share in the pre-Basel II period, the risk-weight increases brought about by Basel II did not affect overall firm-country-product level exports for destinations whose banks are on average rated A1-Baa3 rated.

In Table 8 column 2, we find similar results for the non-OECD sample. Here the tests rely on the triple interactions $D\_Aaa\text{-}Aa3_c \times D\_BASELII_t \times S_{f,c,p,t=1}$ and $D\_A1\text{-}Baa3_c \times D\_BASELII_t \times S_{f,c,p,t=1}$, for which the coefficient estimates are 0.0789 and 0.0747. While these coefficient estimates have the correct signs (for high letter of credit-based export shares in the pre-Basel II, the drop in risk weights is associated with positive total trade growth) neither of them are statistically significant.

The failure to find a statistically significant result would suggest that Basel II did not affect total trade at the firm-product-country level. These growth regression results are consistent with accommodation through payment terms following that Basel II instigated changes in risk-weights.

Importantly, in contrast to concerns expressed by officials of international bodies, evidence from the Turkish exports data suggests that Basel II does not appear to have affected overall trade flows.

In the next section we conduct a series of regressions using our baseline models to check the robustness of our main results.

4. Robustness checks

To rule out alternative explanations of our results, we re-estimate variations of our baseline regression models first using the OECD and then the non-OECD samples. In a third step, we address econometric issues that could potentially affect our results.

4.1. OECD Sample

Our robustness checks for the OECD sample are presented in Table 9. In the first column we apply a version of the so-called "mid-point" regressions proposed by Davis and Haltiwanger (1992).[35] Our endogenous variable, the change in the share of letter of credit-based exports of a firm to a country for a product over total exports of the same firm with the same country and the same product, results in a large number of zeros, which could potentially affect our estimates. To implement a version of the said "mid-point" regressions, we first re-calculate the letter of credit-based export shares as follows:

$$S_{f,c,p,t} = EXPORTS\_LC_{f,c,p,t} / [(EXPORTS\_TOTAL_{f,c,p,t} + EXPORTS\_TOTAL_{f,c,p,t-1}) \times ½]$$
$$S_{f,c,p,t-1} = EXPORTS\_LC_{f,c,p,t-1} / [EXPORTS\_TOTAL_{f,c,p,t} + EXPORTS\_TOTAL_{f,c,p,t-1}) \times ½]$$

Then, we first-difference these shares as before (i.e., $\Delta S_{f,c,p,t} = S_{f,c,p,t} - S_{f,c,p,t-1}$) and re-estimate equations (1) through (3) as before. In Table 9 column 1, the mid-point regression coefficient estimate for the *D_A1-Baa3×D_BASELII* interaction is equal to -0.00180, which is statistically significant at the 1%-level. This coefficient estimate implies a -3.96% (=-0.00180/0.0455) decrease in the letter of credit-based export shares for the A1-Baa3 rating range compared to the pre-Basel II level, and a risk-weight elasticity of letter of credit-based export shares of -0.0264 (=[-0.00180/0.0455]/[(0.50-0.20)/0.20]).

In Table 9 columns 2 and 3, we check whether our results could be due to demand factors that our fixed-effects cannot fully account for.[36] In column 2, we add two additional variables to better control for demand our baseline model of Eq. (1): $\Delta IMPORTS_{c,t}$ are the global imports of country *c* in annual period *t* taken from the IMF DOTS database, excluding any exports from Turkey to that destination, whereas $\Delta SOVEREIGN\_RATING_{c,t}$ is equal to changes in sovereign rating for country *c* during period *t* (from Moody's, S&P and Fitch). Adding these two variables does not materially affect our results: in Table 9, column 2 the coefficient estimate *D_A1-Baa3×D_BASELII* is equal to -0.00570

[35] See also Davis, Haltiwanger and Schuh (1998). We thank an anonymous referee for suggesting this exercise.
[36] We cannot add country×time fixed-effects in our regressions, as these would fully account for the interactions (for ex., *D_A1-Baa3×D_BASELII* for the OECD sample) with which we conduct our tests.

and statistically significant at the 1%-level, which is comparable to the baseline regression coefficient estimate of -0.00537 in column 1 of Table 5. In column 3, we replace $\Delta IMPORTS_{c,t}$ and $\Delta SOVEREIGN\_RATING_{c,t}$ with a country-level time trend: the coefficient estimate for *D_A1-Baa3×D_BASELII* is now equal to -0.00545 and statistically significant at the 1%-level, i.e., little affected. We conclude that our results are not driven by country-level demand factors that remain unaccounted for.

Another potential concern is whether our results could be driven by possible (but unlikely) shifts across credit conversion factors. As mentioned in the Introduction, under both Basel I and Basel II, credit conversion factors remain constant at 20%, 50% or 100% given the type of letter of credit, which might depend on firm, export destination country, and/or product category specifics (see Appendix Table A1).[37] While the TSI dataset does not allow us to observe the types of letters of credit used in Turkish exports (and hence the credit conversion factors that apply), the first-difference of the dependent variable would remove any related effects at the firm-country-product level so long as they remain constant over two consecutive annual periods. That said, first-differencing cannot control for shifts from one credit conversion factor category into another, potentially in response to Basel II adoption. Our results are unlikely to be driven by such change, because they would need to be systematic, which is unlikely. The letters of credit held by Turkish banks are issued by the foreign importer's bank: it is highly improbable that counterparty banks in 25 other OECD countries would *systematically* adjust after Basel II the types of letter of credit that they issue to their importer clients in order to accommodate the Turkish exporters or their Turkish banks. This said, we nevertheless check whether our results stand when we limit the sample to the particular case of metals. This group of products mostly includes standardized manufactured goods (such as sheet and rod iron and steel) that can be easily collateralized, in which case the credit-conversion factor is equal to 20%. Unsurprisingly, given the possibility of collateralization for this group of products, the average share of letter of credit-based exports is higher for metals: 0.3689 in the pre-Basel II period for exports to countries whose banks are rated A1-Baa3 on average. Accordingly, we observe a stronger effect: in column 4 of Table 7, the coefficient estimate for β_1 equals -0.0749, which is statistically significant at the 10%-level. The corresponding percentage change in exports countries with A1-Baa3 rated banks is equal to a drop of 20.30%. As a result, the HS2 product category for metals exhibits a larger risk-weight elasticity of letter of credit-based exports: -0.1354 (=-0.0749/1.50). A 10% increase in the risk-weight of letters of credit leads to a 1.4% drop in the share of letter of credit-financed exports of metal (sheet and rod iron and steel) products to countries whose banks are, on average, rated in the A1-Baa3 range. We conclude that our results cannot be driven by credit conversion factor related effects.

[37] It should be noted that (i) non-binding letters of credit (with a 0% credit conversion factor) that are also listed in Appendix Table A1 have little use in international trade as they can be revoked when the exporter's bank requests payment; and (ii) trade related letters of credit are typically reported to have three to four months of maturity on average, i.e., involve shorter horizons than the one-year of maturity mentioned in Appendix Table A1 (e.g., ICC, March 31, 2009 and October 26, 2011; and SWIFT October, 2009).

Finally, in the last column of Table 9, we conduct a placebo test by using the fictitious Basel II adoption date of July 1st, 2011 after limiting our sample to July 2009-June 2012 period. One could argue that changes implied by Basel II, including those involving risk-weights for the letters of credit, were largely anticipated prior to adoption. This could be because Turkish banks knew in detail how the new regulation was likely to affect different aspects of their operations thanks to Basel II Impact Studies that they were asked to conduct by the banking regulator. As a result, banks might have started to implement changes to their various policies prior to July 1, 2012 in anticipation of the new set of rules to be imposed. In the placebo test regressions presented in the last column of Table 9, the coefficient estimate for *D_A1-Baa3&NR×D_BASELII* is positive (i.e., of the opposite sign than the one expected) and not statistically significant. We conclude that there are no discernable effects that could be due to banks' anticipation of Basel II-related changes to letter of credit risk-weights. [38]

4.2. Non-OECD sample

In this next section, we repeat the above listed robustness checks with the non-OECD sample using our baseline regression Eq. (2). We first assume that the letters of credit have, on average, remaining maturities *longer* than three months. The results are presented in Table 10.

In column 1 of Table 10, the mid-point regressions indicate that both *D_Aaa-Aa3×D_BASELII* and *D_A1-Baa3&NR×D_BASELII* interactions are positive and statistically significant at the conventional levels. In fact, the coefficient estimate for *D_Aaa-Aa3×D_BASELII* is equal to 0.00354 (statistically significant at the 5%-level) is larger than that for *D_A1-Baa3&NR×D_BASELII*, which is equal to 0.00277 (statistically significant at the 1%-level). This makes sense as the risk-weight change for Aaa-Aa3 rating range (from 100% down to 20%) is larger than that for A1-Baa3 range plus the non-rated non-OECD country banks (from 100% to 50%). However, in an (unreported) one-sided t-tests, the null hypothesis for the equality of these two coefficient estimates cannot be rejected. As in the case of OECD mid-point regressions, these non-OECD coefficient estimates imply smaller changes, due to the way in which we calculate the "mid-point" export shares, which now take into account extensive as well as intensive margins.

In columns 2 and 3 of Table 10, our coefficient estimates of interest are little affected by the addition of additional controls for country-level demand effects. The coefficient estimates for the *D_Aaa-Aa3×D_BASELII* interaction in columns 2 and 3 of Table 10 still retain the correct sign, but remain statistically insignificant, whereas those for *D_A1-Baa3&NR×D_BASELII* are equal to 0.00448 and 0.00407 and statistically significant at the 5%-level. These results are very similar to our main results for the OECD sample, i.e., those of column 1 of Panel A in Table 6.

[38] The placebo tests also suggest that the triple-difference regressions remove the differences in pre-Basel II time trends between the treatment and control groups that we observe in the non-differenced dependent variable.

In Table 10, column 4 we restrict ourselves to metals products in order to focus to a HS2 product category for which credit conversion is highly likely to be 20% (due to the fact that the underlying rod or sheet metal can be easily collateralized). The *D_Aaa-Aa3×D_BASELII* interaction has a positive but statistically insignificant coefficient estimate, whereas *D_A1-Baa3&NR× D_BASELII* interaction has a coefficient estimate of 0.0587 that is statistically significant the 5%-level. Given that the average share of letter of credit-financed exports to non-OECD countries whose banks are rated A1 to Baa3 on average is equal to 0.1708 in the year prior to Basel II adoption, the coefficient estimate of 0.0587 suggests an increase of 34.37% in share of exports in response to a 50% decrease in risk-weights (from 100% to 50%). The corresponding risk-weight elasticity of share of exports is equal to -0.6874: a 10% decrease in the risk-weight leads to a 6.87% increase in the share of exports when metals are concerned. This result further suggests that our results are not likely to be related with changes in the mix of credit conversion factors around Basel II adoption date of July 1, 2012.

In the last column of Table 10, we conduct a placebo test assuming a fictitious Basel II adoption date of July 1, 2011 and using July 1, 2010 – June 30, 2012 data. The coefficient estimates for *D_Aaa-Aa3×D_BASELII* and *D_A1-Baa3&NR×D_BASELII* are equal to 0.00275 and 0.0000729, respectively, but are not statistically significant.

We also repeat the same robustness checks under the assumption that the letters of credit issued by non-OECD country banks have *shorter* than three months of remaining maturity, but we do not comment on them here for the sake of brevity. The related results, which are presented in Appendix Table A2, are very much in line with those of Table 10.

Next, we conduct additional robustness checks to address potential econometric issues, and comment on other potential concerns.

4.3. Other concerns and further robustness checks

In this section, first we provide formal statistical tests of differences in trends in the levels of letter of credit-based export shares suggested by Fig. 2.[39] To do so, we re-run the cross-sectional versions of our baseline equations (1) through (3) with one year of pre-Basel II data, after replacing country×product and product×time fixed-effects with product fixed-effects.[40] In columns 1, 3, and 5 of Appendix Table A3, which correspond to OECD and non-OECD samples with different assumptions about letters of credit maturities, the coefficient estimates of *D_BASELII* interactions in these cross-sectional regressions are all statistically significant at the conventional levels: there are non-parallel trends in the pre-Basel II data between the treated and control groups. In further tests, we check whether these differences disappear once we further difference our dependent variable. This is indeed the case: in columns 2, 4, and 6 of Appendix Table A3, the coefficient estimates of *D_BASELII* interactions are

[39] We thank an anonymous referee for suggesting these tests.
[40] If we do not replace country×product by product fixed-effects we would not be available to estimate *D_BASELII* related interactions, which would be fully explained by country×product interactions.

no longer statistically in new cross-sectional regressions with double-differenced pre-Basel II data. These cross-sectional regressions justify our use of Khandelwal, Schott and Wei (2013) approach described above.

Second, we check whether the statistical significance of our results are dependent on our assumption regarding the clustering of the standard errors (along country-time dimensions). To do so, we re-estimate regression equations (1) through (3) under other plausible clustering options and present the related t-statistics under the coefficient estimates in Appendix Table A4:[41] firm-level clustering (related t-statistics are in square-brackets); firm and country×time two-way (i.e., double) clustering (t-statistics in curly-brackets); firm, product and country×time three-way (i.e., triple) clustering (t-statistics within angle brackets). The coefficient estimates for the *D_BASELII* interactions remain statistically significant when we implement these different types of clustering.

Third, we discuss whether there could be alternative explanations for our results and to what extent the underlying scenarios that might be driving them are economically plausible. One potential concern is whether the country-level average bank-rating proxy we use for the rating of individual counterparty-banks' ratings (because the identities of the latter are not observable in the data) may generate a bias in our estimates. Our identification relies on Basel II triggered risk-weights changes assuming constant bank ratings. A bias could arise if bank ratings were to display significant heterogeneity for a given country across time. Despite the fact that we constrain country-level bank rating's average to remain within a rating range that corresponds to a risk-weight category as defined by Basel II (e.g., Aaa-Aa3 or A1-Baa3), such heterogeneity might possibly hide a shift in the agency ratings of letter of credit issuing banks. If they were to occur, such shifts are unlikely to be driving our results: for our results to be explained by shifts within a rating-range, there would have to be a *worsening* of ratings for letter of credit issuing 25 OECD-member countries' banks and a contemporaneous *improvement* of ratings for 133 non-OECD-member countries' institutions, just around the time of Basel II adoption in Turkey. Nevertheless, to do away with this concern, we examine the distribution of bank ratings for these two sets of countries over time to see whether there is any large heterogeneity or major shifts in bank ratings. First, in Appendix Table A5, we tabulate, by rating ranges that correspond to risk-weight categories, the means of our proxy for country-level weighted-average counterparty bank rating in the annual periods that precede and follow Basel II adoption of July 1, 2012.[42] For the OECD sample, the means of the Aaa-Aa3 range in the pre- versus post-Basel II counterparty bank ratings remain equivalent to a rating of Aa3, and the observed increase from 4.20 to 4.35 is not statistically significant in a t-test. In contrast, the counterparty rating for the A1-Baa3 rating-range increases by 0.60, which is statistically significant at the 5%-level in a t-test. Even if the latter difference corresponds to a worsening

[41] In Appendix Table A4, column 1 corresponds to column 1 of Table 5, column 2 to column 1 in Panel A of Table 6, and column 3 to column 1 in Panel B of Table 6.
[42] To do so we start from the highest rating and give a numerical value of 1 to the rating of Aaa, 2 to Aa1, 3 to Aa2, 4 to Aa3, 5 to A1, and so on.

of average rating for the group for which the risk weight increases from 0.20 to 1.00, the change is only from 6.49 (i.e., A2) to 7.09 (A3): this observed change is in the higher half of the A1-Baa3 rating-range. In Appendix Table A5, when we repeat the exercise with the non-OECD sample, none of the means that correspond to ratings ranges Aaa-Aa3, A1-Baa3, and Ba1-B3 show important variation between pre- and post-Basel II periods (none of the t-tests for the significance of the means is statistically significant). So our country-level proxy for counterparty bank ratings show little variation over time. In a second step, to further investigate the issue, we calculate the standard deviations of bank ratings individually for each of the 25 OECD and 133 non-OECD countries separately for the annual periods t=1 (July 1, 2011 through June 30, 2012) and t=2 (July 1, 2012 through June 30, 2013). Then, we calculate the bank total assets-weighted standard deviations (given that this is also the weighting scheme used for the country-level ratings proxy) after dropping countries for which we only have a single rated-bank. We allow for the country-annual period level standard deviations to be equal to zero if all banks in that country have the same rating in that annual period. For the OECD, the weighted-average standard deviation of country-level bank ratings is equal to 1.428 in t=1 and 1.447 in t=2. The difference of 0.019 is not statistically significant in a t-test. For the non-OECD sample, the weighted-average standard deviation of country-level bank ratings is equal to 1.044 in t=1 and 1.077 in t=2. Again, the difference of 0.033 is not statistically significant in a t-test. In a visual test, we also plot the country-level first-difference (i.e., t=2 minus t=1) of weighted-average bank ratings on the first-difference of associated standard deviations and present the resulting graphs in Figure 3. Large heterogeneity in country-level bank ratings would suggest a dispersed plot changes in mean ratings on changes in their standard deviations. In Figures 3.a (for the OECD sample) and 3.b (for the non-OECD sample) we observe the opposite: the data points of the scatter plots are highly localized around the origin. This suggests that our country-level proxy for counterparty bank ratings shows very little variation over time both in its mean and standard deviation. We conclude that Figure 3 and the t-tests mentioned above rule out the possibility that our results are driven by shifts or heterogeneity in bank ratings.

Finally, it could be argued that, rather than risk-weight changes, our results could be driven by other factors, such as importers' move towards higher-rated letter of credit-issuing banks in the export destination countries, a shortening of the maturities of the letters of credit, or our specifications not fully accounting for (through firm×time fixed-effects) changes in exporters' Turkish bank relations. Our findings provide indirect answers to these concerns. For example, our results imply that importing firms did not freely substitute away from lower rating banks, possibly due to the associated switching costs: Niepmann and Schmidt-Eisenlohr (2015) note that the letter of credit business is firm- and country-specific and as a result concentrated (given the country in question). Of course, there may be some cases that fit one or more of the above-mentioned concerns. But, given the granularity of our data, we find it hard to contemplate scenarios under which any remaining unexplained variation to be systematic enough over large numbers of firms to drive our results, and this, in opposite directions for the OECD

and non-OECD samples. We conclude that our results are robust and cannot be explained by alternative stories.

5. Conclusion

In this paper we provide a first test of whether risk-based capital standards affect firm-level international trade patterns after controlling for confounding effects at the firm-export destination country-product code-level. In the Turkish case that we study, all banks located in Turkey were required to adopt Basel II in its Standardized Approach version by July 1, 2012. This has led to changes in the risk-weights that banks need to apply to export-related letters of credit that they hold. Before July 1, 2012, under Basel I, the risk-weight was fixed given the country of origin (OECD-member or otherwise) of the foreign counterparty bank that issues the letter of credit. With Basel II, Turkish banks had to apply risk-weights as a function of the agency rating of the counterparty bank (while the related credit conversion factors remained unaffected). These changes generate two sets of identification schemes that we take to two different samples of Turkish exports to the OECD- and non-OECD-member countries. We estimate difference-in-differences regressions that implicitly take into account firm-country-product confounding factors (through first-differencing of the dependent variable).

We find that following increases (decreases) in the risk-weights for certain counterparty rating categories (A1-Baa3), the share of letter of credit-based exports to the associated OECD (non-OECD) countries decrease (increase) when the Turkish banks' capital charges for carrying these instruments as off-balance sheet items are lowered (increased). These results are robust to the addition of various fixed-effects and control variables. They also suggest that the impact of Basel II adoption had opposite effects on Turkish exports destined to OECD- versus non-OECD-member countries that, otherwise, have banks with similar agency ratings on average. However, while the letter of credit-based export shares are affected by Basel II adoption in Turkey, the total exports (at the firm-product-country level) are not: in our tests, we cannot detect growth in firm-product-country level exports when all export financing methods are combined. This suggests that the effect of higher risk-based capital requirements on exports is subtler than the one anticipated in the popular press and by international trade organizations.[43] In fact, our results indicate that, in the particular case of letters of credit financed exports, the shocks to the trade-finance channel appear to play a less important role in determining the export flows, than the one suggested by the findings of Amiti and Weinstein (2011) or Chor and Manova (2012). Our findings are more in line with those of Eaton et al. (2011) who note that the bulk of the decrease in international trade during the Great Recession could be attributed to a fall in durable goods' trade (probably linked with a fall in demand for such goods) rather than a pure finance-channel effect. After controlling for

[43] In a similar vein, Fraisse, Lé, and Thesmar (2015) find that the 2008 Basel II adoption in France had subtler effects than expected. Under Basel II aggregate firm borrowing and investment increased in France, allowing the preservation of jobs during a crisis period as average bank capital required for industrial loans decreased by 2% on average.

potential confounding effects (including country- or product-demand) at the firm-product-country-level, we find that while letter of credit based export shares are affected, the growth of exports (after combining all financing methods) are not.

As such, our results also contribute to the policy debate regarding the risk that letters of credit used in international trade present for banks. There is a concern that the higher-than-warranted credit conversion factors and/or risk-weights for letters of credit can stifle international trade. This is because these trade instruments appear to be of much lower risk than regular lending to the real sector.[44] Some of these concerns are already incorporated in the Basel standards. For example, in 2011, for banks confirming export-related letters of credit, the Basel Committee waived the so-called "sovereign floor" applicable under the Standardized Approach.[45] Similarly, for the calculation of the capital requirements and liquidity risk, Basel III now stipulates that the credit conversion factors for most categories of letters of credit should be equal to 20%, effectively reducing them in the majority of cases. Our findings suggest that these recent changes are likely to have a positive impact on international trade through channels similar to the ones discussed in our paper.

[44] A finance loss register in 2013 based on 15 million trade transactions showed that trade finance had a 0.021% default rate and a high recovery rate of trade finance products. This was three times lower than Moody's yearly A1 customer default rate of 0.065% (Trade Finance, 2013).

[45] The sovereign floor refers to the fact that the risk-weight of the counterparty bank issuing a letter of credit cannot be lower than the risk-weight associated with the rating of the sovereign in which it is located.

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Table 1
Basel I and Basel II risk-weights applied to off-balance sheet commercial letters of credit

Risk-Weights based on long-term agency ratings					
Basel I		Basel II			
Risk-Weights		Risk-Weights		Agency Rating Categories	
1	2	3	4	5	6
OECD	Non-OECD	letter of credit maturity < 3 months (in case original short-term ratings do not exist)	letter of credit maturity > 3 months	Moody's	Fitch or S&P
0.20	1.00	0.20	0.20	Aaa to Aa3	AAA to AA-
			0.50	A1 to Baa3	A+ to BBB-
		0.50	1.00	Ba1 to B3	BB+ to B-
		1.50	1.50	Caa1 and below	CCC+ and below
		0.20	0.50	Non-rated (NR)	Non-rated (NR)

This table presents the Basel I and Basel II risk-weights applied to foreign bank liabilities held by the Turkish banks for a given credit rating, including confirmed export related commercial letters of credit issued by banking institutions domiciled in other countries.

Table 2
Summary statistics for the dependent variable

Panel A

	Letter of credit-based share of exports at the firm-country-product level by annual period					
	OECD Sample			Non-OECD Sample		
	7.1.2011 – 6.30.2012	7.1.2012 – 6.30.2013	t-stat	7.1.2011 – 6.30.2012	7.1.2012 – 6.30.2013	t-stat
	0.0363	0.0343	1.83 *	0.0584	0.0550	2.79 ***
	(0.1633)	(0.1614)		(0.2147)	(0.2101)	
	[47,183]	[47,183]		[63,040]	[63,040]	

Panel B

	Letter of credit-based share of exports at the firm-country-product level by annual period and counterparty rating ranges that correspond to risk-weight categories					
	OECD Sample			Non-OECD Sample		
Counterparty bank ratings	7.1.2011 – 6.30.2012	7.1.2012 – 6.30.2013	t-stat	7.1.2011 – 6.30.2012	7.1.2012 – 6.30.2013	t-stat
Aaa-Aa3	0.0130	0.0118	0.97	0.0958	0.0896	0.60
	(0.1010)	(0.0970)		(0.2594)	(0.2520)	
	[13,375]	[13,375]		[1,193]	[1,193]	
A1-Baa3	0.0455	0.0433	1.61 †	0.0729	0.0699	1.70 *
	(0.1813)	(0.1799)		(0.2392)	(0.2366)	
	[33,808]	[33,808]		[37,851]	[37,851]	
Ba1-B3				0.0337	0.0299	2.63 ***
				(0.1629)	(0.1532)	
				[23,996]	[23,996]	

Panel A presents the means for the dependent variable in the annual periods around Basel II adoption date of July 1, 2012. The standard deviations are within parentheses, the number of observations within brackets. Panel B presents the same statistics by rating categories that correspond to different risk-weights, some of which are subject to change with Basel II adoption and drive our identification schemes. T-statistics are provided for two-sided tests, which allow for unequal variances, of the equality of the mean of letter of credit based share of exports. ***, **, * and † denote statistical significance at the 1%, 5%, 10%, and 11% levels, respectively.

Table 3
Additional sample statistics: Industry breakdown

Industry name	HS2 code	Letter of credit-based exports (in %)	
		July 2011 - June 2012	July 2012 - June 2013
Textiles	50-63	19.80	19.71
Metals	72-83	17.41	15.98
Machinery / electrical	84-85	16.48	15.52
Transportation	86-89	13.35	12.25
Stone / glass	68-71	9.63	12.79
Plastics / rubbers	39-40	5.84	5.60
Food	16-24	4.79	4.75
Chemicals and allied industries	28-38	4.16	4.29
Mineral products	25-27	2.96	3.19
Miscellaneous	90-97	2.93	3.44
Wood and wood products	44-49	1.57	1.36
Raw hides, skins, leather and furs	41-43	0.65	0.62
Footwear / headgear	64-67	0.42	0.49

This table presents the letter of credit-based exports by industry for the two annual periods around Basel II adoption date of July 1, 2012. Statistics based on the value of exports at the two-digit Harmonized System (HS2) industry level.

Table 4
Additional sample statistics: Firm, country, and product breakdown

	Panel A. OECD sample							
	Letter of credit financed exports				Non-letter of credit financed exports			
Total number of …	N	Mean	Std. dev.	Median	N	Mean	Std. dev.	Median
firms per country	25	66.28	90.83	39	25	935	924.28	645
products per firm	996	2.38	3.64	1	8,900	2.47	4.02	1
products per country	25	72.76	83.22	37	25	526.64	355.17	445
firms per country-product pair	1,819	1.86	3.05	1	13,166	3.33	7.77	1
products per firm-country pair	1,657	2.04	2.73	1	23,375	1.87	2.65	1
countries per firm-product pair	2,371	1.43	1.12	1	21,962	1.99	2.20	1

	Panel B. Non-OECD sample							
	Letter of credit financed exports				Non-letter of credit financed exports			
Total number of …	N	Mean	Std. dev.	Median	N	Mean	Std. dev.	Median
firms per country	105	32.03	63.97	13	133	214.23	430.44	41
products per firm	1,746	1.88	2.28	1	11,009	2.93	8.05	1
products per country	105	36.28	64.05	16	133	169.43	267.73	48
firms per country-product pair	3,809	1.47	2.30	1	22,534	2.55	4.43	1
products per firm-country pair	3,363	1.66	1.73	1	28,492	2.02	5.03	1
countries per firm-product pair	3,286	1.70	2.29	1	32,240	1.78	2.38	1

This table presents statistics on the various dimensions of the data. N is the number of observations. Statistics based on data from July 1, 2011 through June 30, 2012, i.e., for the annual period prior to Basel II adoption.

Table 5
OECD sample regressions (assuming average remaining letter of credit maturities longer than three months)

	Baseline regression		Additional firm×time fixed-effects		Additional firm×product×time fixed-effects	
	1		2		3	
$D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t$	-0.00537	***	-0.00588	***	-0.00592	***
	(-4.85)		(-3.80)		(-3.79)	
Number of observations	94,366		94,366		94,366	
Regression R^2	0.147		0.343		0.359	
Fixed-effects	c×p, p×t		c×p, p×t, f×t		c×p, f×p_{HS2}×t	
Clustered standard errors	c×t		c×t		c×t	

Assuming average remaining maturities of more than three months for the letters of credit, this table presents the estimates of regression:

$$\Delta S_{f,c,p,t} = \beta_1 \, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$$

in which Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1, 2\}$, denotes annual periods in the two-period panel with first-differenced dependent variable; t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013, i.e., the year that follows Basel II adoption date of July 1, 2012 (t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_A1\text{-}Baa3\&NR_c$ is equal to one if the letter of credit-issuer counterparty-banks in the destination OECD country c have, on average, a long-term credit rating between A1 to Baa3 according to Moody's (A+ to BBB- according to S&P or Fitch) or are non-rated, and for which the letter of credit-associated risk-weight increases from 20% to 50% with Basel II, or zero otherwise; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise; $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product×time (p×t) fixed-effects; and ε is the error term of the OLS regression. All product fixed effects (p) are at the 6-digit Harmonized System (HS6) level, with the exception of the firm-product-time (f×p_{HS2}×t) fixed effects in column 3, which are at the HS2-level. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 6
Non-OECD sample regressions

	Baseline regression		Additional firm×time fixed-effects		Additional firm×product×time fixed-effects	
Panel A	1		2		3	
$D\_Aaa\text{-}Aa3_c \times D\_BASELII_t$	0.00140		-0.00552		-0.00672	
	(0.45)		(-0.62)		(-0.79)	
$D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t$	0.00447	**	0.00530	**	0.00519	*
	(2.52)		(2.15)		(1.78)	
Number of observations	126,080		126,080		126,080	
Regression R^2	0.209		0.406		0.437	
Fixed-effects	c×p, p×t		c×p, p×t, f×t		c×p, f×p(HS2)×t	
Clustered standard errors	c×t		c×t		c×t	
Panel B						
$D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t$	0.00439	**	0.00497	**	0.00482	*
	(2.49)		(2.04)		(1.67)	
Number of observations	126,080		126,080		126,080	
Regression R^2	0.209		0.406		0.437	
Fixed-effects	c×p, p×t		c×p, p×t, f×t		c×p, f×p_{HS2}×t	
Clustered standard errors	c×t		c×t		c×t	

Panel A presents the estimates of regression, which assumes that the letters of credit have average remaining maturities of *more* than three months:

$$\Delta S_{f,c,p,t} = \beta_1\, D\_Aaa\text{-}Aa3_c \times D\_BASELII_t + \beta_2\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$$

Panel B presents the estimates of regression, which assumes that the letters of credit have average remaining maturities of *less* than three months:

$$\Delta S_{f,c,p,t} = \beta_1\, D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$$

in which Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1, 2\}$, denotes annual periods in the two-period panel with first-differenced dependent variable; t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013, i.e., the year that follows Basel II adoption date of July 1, 2012 (t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_Aaa\text{-}Aa3_c$ ($D\_A1\text{-}Baa3\&NR_c$) [$D\_Aaa\text{-}Baa3\&NR_c$] is equal to one if the letter of credit-issuing counterparty-banks in the destination non-OECD country c are, on average, rated between Aaa and Aa3 (A1 through Baa3 or non-rated) [A1 through Baa3 or non-rated] from Moody's throughout the sample period; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise; $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product×time (p×t) fixed-effects; and ε is the error term of the OLS regression. All product fixed effects (p) are at the 6-digit Harmonized System (HS6) level with the exception of the triple firm×product×time (f×p_{HS2}×t) fixed effects in column 3, which are at the HS2-level. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 7
Pooled (OECD plus non-OECD) sample regressions

	Baseline regression		Additional firm×time fixed-effects		Additional firm×product×time fixed-effects	
	1		2		3	
$D\_RW\_INCREASE_c \times D\_BASELII_t$	-0.00512	***	-0.00407	*	-0.00448	***
	(-4.21)		(-1.93)		(-2.59)	
$D\_RW\_DECREASE_c \times D\_BASELII_t$	0.00457	**	0.00551	**	0.00539	**
	(2.55)		(2.36)		(2.05)	
$D\_OECD_c \times D\_BASELII_t$	0.00771	***	0.0114	***	0.0121	***
	(4.37)		(4.24)		(4.44)	
Number of observations	220,446		220,446		220,446	
Regression R^2	0.176		0.334		0.364	
Fixed-effects	c×p, p×t		c×p, p×t, f×t		c×p, f×p_{HS2}×t	
Clustered std. errors	c×t		c×t		c×t	
H_0: $D\_RW\_INCREASE_c \times D\_BASELII_t = D\_RW\_DECREASE_c \times D\_BASELII_t$	18.89	***	10.34	***	10.92	***
H_0: $\lvert D\_RW\_INCREASE_c \times D\_BASELII_t \rvert = \lvert D\_RW\_DECREASE_c \times D\_BASELII_t \rvert$	0.0678		0.189		0.0761	

This table presents the estimates of regression equation:

$$\Delta S_{f,c,p,t} = \beta_1\, D\_RW\_INCREASE_c \times D\_BASELII_t + \beta_2\, D\_RW\_DECREASE_c \times D\_BASELII_t + \beta_3\, D\_OECD_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$$

in which Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t denotes annual periods in the two-period panel of first-differenced dependent variable with $t \in \{1, 2\}$ (t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013; t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_RW\_INCREASE_c$ ($D\_RW\_DECREASE_c$) is equal to one if the risk-weights for letters of credit increase (decrease) with Basel II adoption, which correspond to OECD (non-OECD) countries; D_OECD_c is equal to one if country c is an OECD-member throughout the sample period; $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product×time (p×t) fixed-effects; and ε is the error term of the OLS regression. All product fixed effects (p) are at the 6-digit Harmonized System (HS6) level with the exception of the triple firm-product-time (f×p_{HS2}×t) fixed effects in column 3, which are at the HS2-level. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 8
Firm-destination-product level trade growth regressions

	OECD sample		Non-OECD sample	
	1		2	
$D\_Aaa\text{-}Aa3_c \times D\_BASELII_t \times S_{f,c,p,t=1}$			0.0789	
			(0.27)	
$D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t \times S_{f,c,p,t=1}$	-0.0557		0.0747	
	(-0.30)		(0.43)	
$D\_Aaa\text{-}Aa3_c \times S_{f,c,p,t=1}$			-0.164	
			(-0.58)	
$D\_A1\text{-}Baa3\&NR_c \times S_{f,c,p,t=1}$	0.00471		-0.0150	
	(0.03)		(-0.10)	
$D\_BASELII_t \times S_{f,c,p,t=1}$	0.383	**	0.233	*
	(2.57)		(1.94)	
$S_{f,c,p,t=1}$	-0.350	***	-0.295	***
	(-3.02)		(-2.98)	
Number of observations	94,366		126,080	
Regression R^2	0.315		0.436	
Fixed-effects	c×p×t		c×p×t	
Clustered standard errors	c×t		c×t	

Column 1 of this table presents the estimates of regression equation for the OECD sample:

$$\Delta ln(EXPORTS\_TOTAL_{f,c,p,t}) = \beta_1\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t \times S_{f,c,p,t=1} + \beta_2\, D\_A1\text{-}Baa3\&NR_c \times S_{f,c,p,t=1} + \beta_3\, D\_BASELII_t \times S_{f,c,p,t=1} + \beta_4\, S_{f,c,p,t=1} + \gamma_{c,p,t} + \varepsilon_{f,c,p,t}$$

Column 2 of this table presents the estimates of regression equation for the non-OECD sample:

$$\Delta ln(EXPORTS\_TOTAL_{f,c,p,t}) = \beta_1\, D\_Aaa\text{-}Aa3_c \times D\_BASELII_t \times S_{f,c,p,t=1} + \beta_2\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t \times S_{f,c,p,t=1} + \beta_3\, D\_Aaa\text{-}Aa3_c \times S_{f,c,p,t=1} + \beta_4\, D\_A1\text{-}Baa3\&NR_c \times S_{f,c,p,t=1} + \beta_5\, D\_BASELII_t \times S_{f,c,p,t=1} + \beta_6\, S_{f,c,p,t=1} + \gamma_{c,p,t} + \varepsilon_{f,c,p,t}$$

in which $ln(EXPORTS\_TOTAL_{f,c,p,t})$ is the natural logarithm of the value of total exports (with all types of export financing combined) of firm f to country c for a given product p;; subscript t , with t ∈ {1, 2}, denotes annual periods in the two-period panel (t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013); $D\_Aaa\text{-}Aa3_c$ ($D\_A1\text{-}Baa3\&NR_c$) is equal to one if the letter of credit-issuing counterparty-banks in the destination non-OECD country c are, on average, rated between Aaa and Aa3 (A1 through Baa3 or non-rated) from Moody's throughout the sample period; $D\_BASELII_t$ is equal to one for t=2 post Basel II, and zero otherwise; $S_{f,c,p,t=1}$ is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in (pre-Basel II) period $t=1$ (pre-Basel II) with respect to total exports of f in p to c during the same period; $\gamma_{c,p,t}$ denotes country×product×time (c×p×t) fixed effects; and ε is the error term of the OLS regression. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 9
Robustness checks: OECD Sample (assuming average remaining letters of credit maturities longer than three months)

	Mid-point regression		Additional controls		Additional time-trend		Metals subsample		Placebo regression
	1		2		3		4		5
$D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t$	-0.00180	***	-0.00570	***	-0.00545	***	-0.0749	*	0.00195
	(-2.66)		(-5.06)		(-3.78)		(-1.67)		(-0.66)
$\Delta IMPORTS_{c,t}$			0.00793						
			(0.44)						
$\Delta SOVEREIGN\_RATING_{c,t}$			0.00104	*					
			(1.92)						
$COUNTRY\_TREND_{c,t}$					-0.0000497				
					(-0.10)				
Number of observations	137,406		94,366		94,366		836		63,274
Regression R^2	0.176		0.147		0.147		0.246		0.189
Fixed-effects	c×p, p×t		c×p, p×t		c×p, p×t		c×p, p×t		c×p, p×t
Clustered standard-errors	c×t		c×t		c×t		c×t		c×t

Columns 1, 4 and 5 of this table presents the estimates of regression: $\Delta S_{f,c,p,t} = \beta_1\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$ in which Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1, 2\}$, denotes annual periods in the two-period panel with first-differenced dependent variable; t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013, i.e., the year that follows Basel II adoption date of July 1, 2012 (t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_A1\text{-}Baa3_c$ is equal to one if the letter of credit-issuer counterparty-banks in the destination OECD country c have, on average, a long-term credit rating between A1 to Baa3 according to Moody's (A+ to BBB- according to S&P or Fitch) and for which the letter of credit-associated risk-weight increases from 20% to 50% with Basel II, or zero otherwise; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise. In columns 2 and 3 additional control variables are added: $\Delta IMPORTS_{c,t}$ is total imports (excluding Turkish exports) of country c in period t; $\Delta SOVEREIGN\_RATING_{c,t}$ equals to change in sovereign rating that country c has experienced in period t; and $COUNTRY\_TREND_{c,t}$ is a country-specific time trend. In all regressions $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product×time (p×t) fixed-effects; and ε is the error term of the OLS regression. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 10
Robustness checks: Non-OECD sample (assuming average remaining letter of credit maturities longer than three months)

	Mid-point regression		Additional controls		Additional time-trend		Metals subsample		Placebo regression
	1		2		3		4		5
$D\_Aaa\text{-}Aa3_c \times D\_BASELII_t$	0.00354	**	0.00148		0.00112		0.0365		0.00275
	(2.49)		(0.46)		(0.35)		(0.62)		(0.25)
$D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t$	0.00277	***	0.00448	**	0.00407	**	0.0587	**	0.0000729
	(2.76)		(2.37)		(2.33)		(2.39)		(0.02)
$\Delta IMPORTS_{c,t}$			0.00397						
			(1.22)						
$\Delta SOVEREIGN\_RATING_{c,t}$			-0.000101						
			(-0.08)						
$COUNTRY\_TREND_{c,t}$					-0.000418				
					(-0.75)				
Number of observations	199,198		124,268		126,080		3,820		72,666
Regression R^2	0.275		0.209		0.209		0.231		0.253
Fixed-effects	c×p, p×t		c×p, p×t		c×p, p×t		c×p, p×t		c×p, p×t
Clustered standard-errors	c×t		c×t		c×t		c×t		c×t

Columns 1, 4 and 5 of this table presents the estimates of regression: $\Delta S_{f,c,p,t} = \beta_1\, D\_A1\text{-}Baa3_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$ in which Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1, 2\}$, denotes annual periods in the two-period panel with first-differenced dependent variable; t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013, i.e., the year that follows Basel II adoption date of July 1, 2012 (t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_A1\text{-}Baa3\&NR_c$ is equal to one if the letter of credit-issuer counterparty-banks in the destination OECD country c have, on average, a long-term credit rating between A1 to Baa3 according to Moody's (A+ to BBB- according to S&P or Fitch) or non-rated, and for which the letter of credit-associated risk-weight increases from 20% to 50% with Basel II, or zero otherwise; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise. In columns 2 and 3 additional control variables are added: $\Delta IMPORTS_{c,t}$ is total imports (excluding Turkish exports) of country c in period t; $\Delta SOVEREIGN\_RATING_{c,t}$ equals to change in sovereign rating that country c has experienced in period t; and $COUNTRY\_TREND_{c,t}$ is a country-specific time trend. In all regressions $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product × time (p×t) fixed-effects; and ε is the error term of the OLS regression. All product fixed effects (p) are at the 6-digit Harmonized System (HS6) level. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Figure 1
Pre-"treatment" (Basel II adoption) distribution of share of letter of credit-based exports at the firm-country-product level

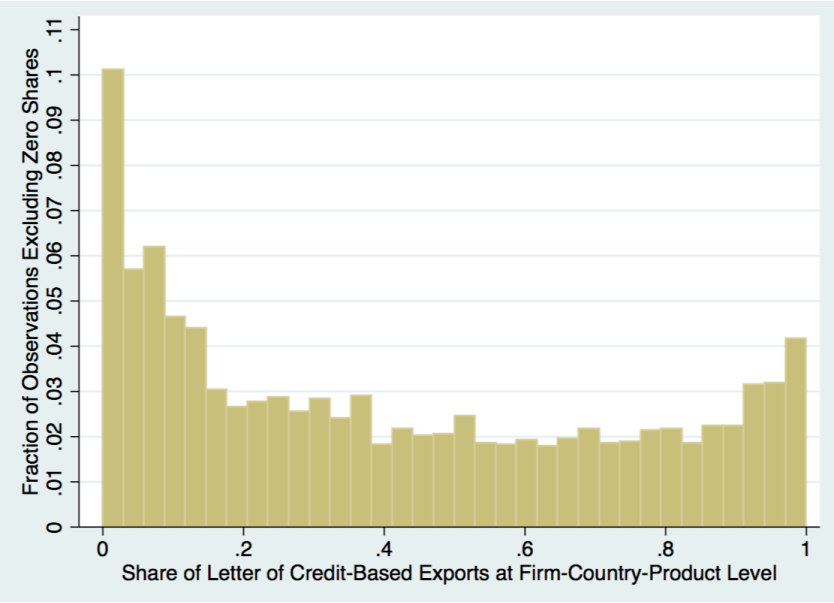

Figure 1.a OECD sample

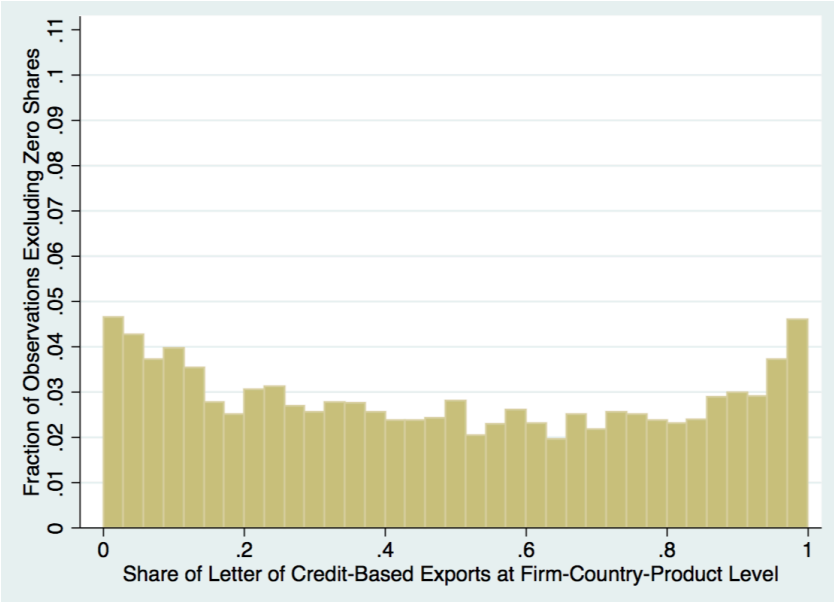

Figure 1.b non-OECD sample

Figure 1.a plots, for the OECD sample, the frequency distribution of the share of letter of credit-based exports at the firm-country-product level in the annual period preceding July 1, 2012 Basel II adoption (i.e., July 1, 2011 – June 30, 2012) after excluding observations for which the share of exports is zero. Figure 1.b does the same for the non-OECD sample.

Figure 2
Pre- and post-"treatment" trends in letter of credit-based share of exports at the firm-country-product level

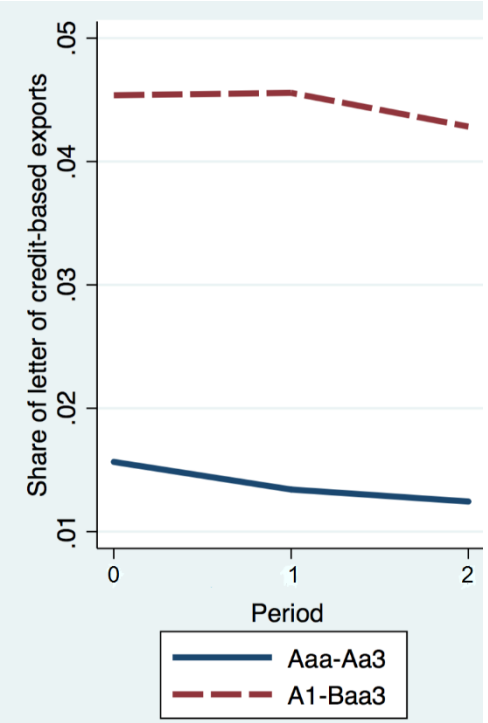

Figure 2.a OECD sample
letter of credit maturity>3 months

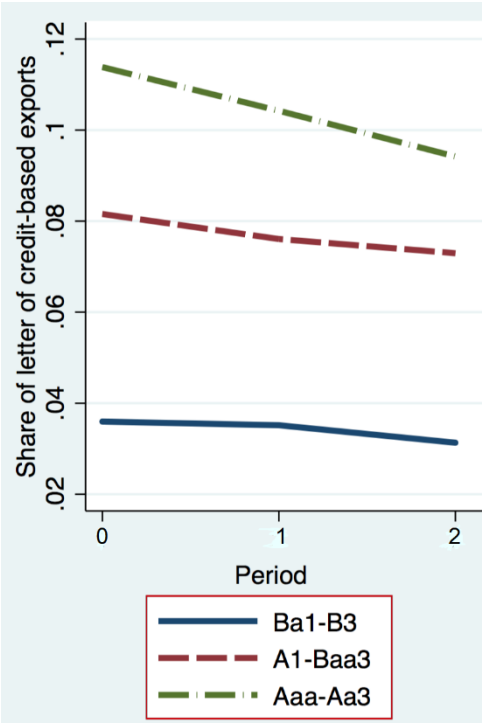

Figure 2.b non-OECD sample
letter of credit maturity>3 months

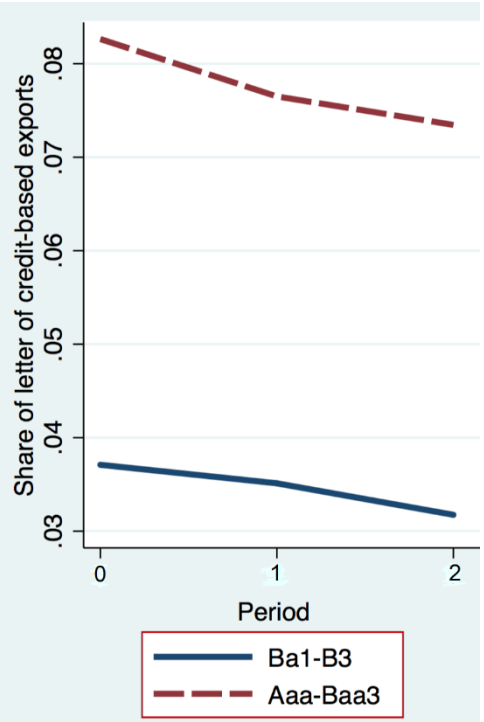

Figure 2.c non-OECD sample
letter of credit maturity<3 months

Figures 2.a, 2.b and 2.c plot, for different annual periods, pre- and post-"treatment" (Basel II adoption) trends in letter of credit-based share of exports at the firm-country-product level for ranges of counterparty bank rating-ranges that correspond to risk-weight categories defined by Basel II (and depicted in Table 1). Pre-Basel II annual periods correspond to t=0 (July 1, 2010-June 30, 2011) and t=1 (July 1, 2011-June 30, 2012), whereas post-Basel II period is t=2 (July 1, 2012-June 30, 2013). Figure 1.a, which corresponds to the OECD sample used in the estimation of Eq.(1), assumes that the letters of credit have a remaining average maturity longer than three months. Figure 1.b, which corresponds to the non-OECD sample used in the estimation of Eq.(2), assumes that the letters of credit have a remaining average maturity longer than three months. Figure 1.c, which corresponds to the non-OECD sample used in the estimation of Eq.(3), assumes that the letters of credit have a remaining average maturity shorter than three months.

Figure 3
Changes in the means and standard deviations of OECD and non-OECD bank ratings

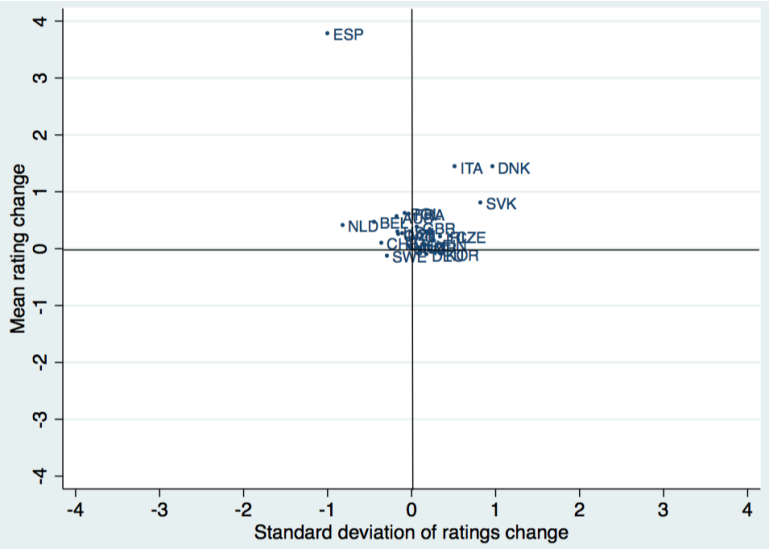

Figure 3.a OECD sample

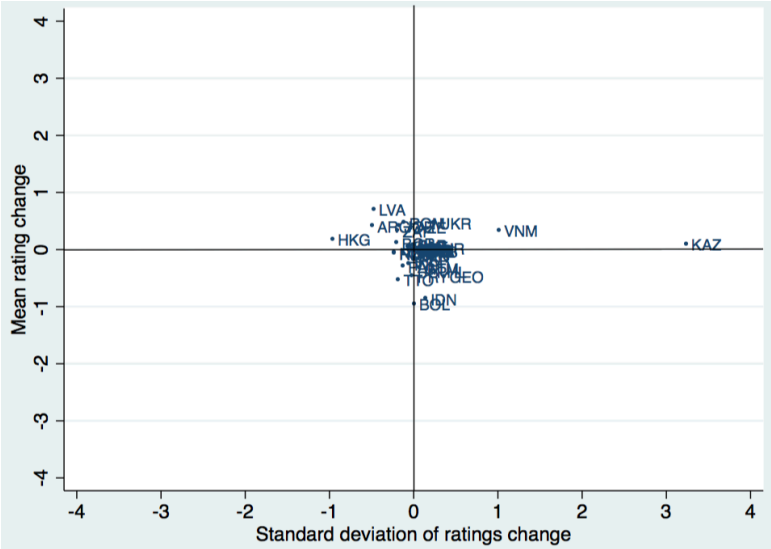

Figure 3.b Non-OECD sample

Figure 3.a plots, the changes in the (total assets weighted) mean of individual bank ratings per OECD country (y-axis) against the changes in the associated standard deviation (x-axis). Figure 3.b plots the same variables for the non-OECD member countries. Agency bank ratings, which are from BankScope database, have been assigned numerical values as follows: 1 to the rating of Aaa, 2 to Aa1, 3 to Aa2, 4 to Aa3, 5 to A1, etc. The country-level bank-total asset weighted means and standard deviations are calculated over two annual periods t=1 (July 1, 2011 through June 30, 2012) and t=2 (July 1, 2012 through June 30, 2013). Changes are calculated between these two annual periods for each country.

Appendix A

Appendix Table A1
Basel I and Basel II credit conversion factors for off-balance sheet commercial letters of credit

BDDK Directive of November 1, 2006 (under Basel I) Article Number	BDDK Directive of June 28, 2012 (for Basel II) Article 5 (2) and Article Number	Credit Conversion Factor	Letter of Credit Type
5 (1) a) 3)	5 (3) a) 3)	100%	Export-related confirmed-letters of credit
5 (1) b) 6)	5 (3) b) 5)	50%	Letters of credit that do not have a CCF of 100%, 20% or 0%.
5 (1) c) 2)	5 (3) c) 2)	20%	Letters of credit with maturity less than one year and in which the exported good serves as collateral.
5 (1) ç) 2)	5 (3) ç) 2)	0%	Non-binding letters of credit that do not require a payment to the recipient [exporter].

This table presents the Basel I and Basel II credit conversion factors (CCFs) for off-balance sheet commercial letters of credit (Letters of credit) held by the Turkish banks after being issued by a foreign counterparty bank for payments to Turkish exporters (BDDK Directives of November 1, 2006 and June 28, 2012).

Appendix Table A2
Robustness checks: Non-OECD sample (assuming an average remaining maturity shorter than three months for the letters of credit)

	Mid-point regression		Additional Controls		Additional time-trend		Metals subsample		Placebo regression
	1		2		3		4		5
$D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t$	0.00279	***	0.00440	**	0.00346	†	0.0581	**	0.000152
	(2.83)		(2.35)		(1.47)		(2.40)		(0.05)
$\Delta IMPORTS_{c,t}$			0.00403						
			(1.24)						
$\Delta SOVEREIGN\_RATING_{c,t}$			-0.000128						
			(-0.10)						
$COUNTRY\_TREND_{c,t}$					0.000633				
					(0.90)				
Number of observations	199,210		124,268		126,080		3,820		72,666
Regression R^2	0.275		0.209		0.209		0.231		0.253
Fixed-effects	c×p, p×t		c×p, p×t		c×p, p×t		c×p, p×t		c×p, p×t
Clustered standard-errors	c×t		c×t		c×t		c×t		c×t

in which Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1, 2\}$, denotes annual periods in the two-period panel with first-differenced dependent variable; t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013, i.e., the year that follows Basel II adoption date of July 1, 2012 (t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_Aaa\text{-}Baa3\&NR_c$ is equal to one if the letter of credit-issuer counterparty-banks in the destination OECD country c have, on average, a long-term credit rating between A1 to Baa3 according to Moody's (A+ to BBB- according to S&P or Fitch) or are non-rated, and for which the letter of credit-associated risk-weight increases from 20% to 50% with Basel II, or zero otherwise; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise; $\Delta SOVEREIGN\_RATING_{c,t}$ is the difference in countries' ratings between t=1 to t=2 (i.e., from July 2010-June 2011 to July 2011-June 2012) after having turned sovereign ratings into numerical values (1 for Aaa, 2 for Aa1, 3 for Aa2, etc.); $COUNTRY\_TREND_{c,t}$ is a country-specific time trend; $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product×time (p×t) fixed-effects; and ε is the error term of the OLS regression. All product fixed effects (p) are at the 6-digit Harmonized System (HS6) level. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Appendix Table A3
Cross-sectional pre-Basel II regressions: Tests of parallel pre-treatment trends

	OECD sample (letter of credit maturity>3-months)			Non-OECD sample (letter of credit maturity>3-months)			Non-OECD sample (letter of credit maturity<3-months)		
Dependent variable	$\Delta S_{f,c,p}$		$\Delta\Delta S_{f,c,p}$	$\Delta S_{f,c,p}$		$\Delta\Delta S_{f,c,p}$	$\Delta S_{f,c,p}$		$\Delta\Delta S_{f,c,p}$
	1		2	3		4	5		6
D_Aaa-Aa3$_c$				-0.00385	*	0.0000729			
				(-1.67)		(0.02)			
D_A1-Baa3&NR$_c$	0.00476	***	0.00195	-0.00494	**	0.00275			
	(3.49)		(0.50)	(-2.48)		(0.20)			
D_Aaa-Baa3&NR$_c$							-0.00388	*	0.000152
							(-1.71)		(0.04)
Number of observations	47,183		31,637	63,040		36,333	63,040		36,333
Regression R^2	0.0560		0.0783	0.0484		0.0613	0.0484		0.0613
Fixed-effects	p		p	p		p	p		p
Clustered standard-errors	c		c	c		c	c		c

This presents *cross-sectional* regressions with *first-differenced* letter-of-credit based export shares:

Column 1: $\Delta S_{f,c,p} = \beta_1\ D\_A1\text{-}Baa3\&NR_c + \gamma_p + \varepsilon_{f,c,p}$
Column 2: $\Delta\Delta S_{f,c,p} = \beta_1\ D\_A1\text{-}Baa3\&NR_c + \gamma_p + \varepsilon_{f,c,p}$
Column 3: $\Delta S_{f,c,p} = \beta_1\ D\_Aaa\text{-}Aa3_c + \beta_2\ D\_A1\text{-}Baa3\&NR_c + \gamma_p + \varepsilon_{f,c,p}$
Column 4: $\Delta\Delta S_{f,c,p} = \beta_1\ D\_Aaa\text{-}Aa3_c + \beta_2\ D\_A1\text{-}Baa3\&NR_c + \gamma_p + \varepsilon_{f,c,p}$
Column 5: $\Delta S_{f,c,p} = \beta_1\ D\_Aaa\text{-}Baa3\&NR_c + \gamma_p + \varepsilon_{f,c,p}$
Column 6: $\Delta\Delta S_{f,c,p} = \beta_1\ D\_Aaa\text{-}Baa3\&NR_c + \gamma_p + \varepsilon_{f,c,p}$

where Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$) and $\Delta\Delta$ indicates double first-differencing (i.e., $\Delta\Delta S_t = \Delta S_t - \Delta S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1\}$, denotes the single cross-sectional and covers July 1, 2011 through June 30, 2012 (t=0 and t=-1, which are implicit in the first- and double-differencing, cover July 1, 2010 through June 30, 2011, and July 1, 2009 through June 30, 2010, respectively); *D_Aaa-Aa3*$_c$ (*D_A1-Baa3&NR*$_c$) [*D_Aaa-Baa3&NR*$_c$] is equal to one if the letter of credit-issuing counterparty-banks in the destination non-OECD country c are, on average, rated between Aaa and Aa3 (A1 through Baa3 or non-rated) [A1 through Baa3 or non-rated] from Moody's throughout the sample period; γ_p denotes product (p) fixed effects; and ε is the error term of the OLS regression. Regressions robust standard errors are clustered at the country (c) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Appendix Table A4
Robustness checks: Baseline regressions with different clustering options for the standard errors

	OECD sample (letter of credit maturity>3-months)		Non-OECD sample (letter of credit maturity>3-months)		Non-OECD sample (letter of credit maturity<3-months)	
	1		2		3	
$D\_Aaa\text{-}Aa3_c \times D\_BASELII_t$			0.00140			
(c×t clustering)			(0.45)			
[f clustering]			[0.13]			
{c×t and f double-clustering}			{0.21}			
⟨c×t and f and p triple-clustering⟩			⟨0.17⟩			
$D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t$	-0.00537		0.00447			
(c×t clustering)	(-4.85)	***	(2.52)	**		
[f clustering]	[-2.31]	**	[1.78]	*		
{c×t and f double-clustering}	{-3.08}	***	{1.96}	**		
⟨c×t, f and p triple-clustering⟩	⟨-2.72⟩	***	⟨1.86⟩	*		
$D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t$					0.00439	
(c×t clustering)					(2.29)	**
[f clustering]					[1.76]	*
{c×t and f double-clustering}					{1.94}	*
⟨c×t, f and p triple-clustering⟩					⟨1.84⟩	*
Number of observations	94,366		126,080		126,080	
Regression R^2	0.147		0.209		0.209	
Fixed-effects	c×p, p×t		c×p, p×t		c×p, p×t	
Clustered standard-errors	c×t		c×t		c×t	

This Appendix table presents the estimates of the baseline regression equations:

Column 1: $\Delta S_{f,c,p,t} = \beta_1\, D\_A1\text{-}Baa3_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$
Column 2: $\Delta S_{f,c,p,t} = \beta_1\, D\_Aaa\text{-}Aa3_c \times D\_BASELII_t + \beta_2\, D\_A1\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$
Column 3: $\Delta S_{f,c,p,t} = \beta_1\, D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t + \gamma_{c,p} + \gamma_{p,t} + \varepsilon_{f,c,p,t}$

in which the standard errors are (i) country×time (c×t) clustered with the corresponding t-statistics presented in parentheses, (ii) firm (f) clustered with the corresponding t-statistics presented in square brackets, (iii) country×time and firm (c×t and f) double-clustered with the corresponding t-

statistics presented in curly brackets, and (iv) country×time, firm and product (c×t, f and p) triple-clustered with the corresponding t-statistics presented in angle brackets. Δ is the first-difference operator (i.e., $\Delta S_t = S_t - S_{t-1}$); S is the share of letter of credit-based exports for firm f exporting product p to export-destination country c in period t with respect to total exports of f in p to c during t; subscript t , with $t \in \{1, 2\}$, denotes annual periods in the two-period panel with first-differenced dependent variable; t=1 covers July 1, 2011 through June 30, 2012; and t=2 corresponds to July 1, 2012 through June 30, 2013, i.e., the year that follows Basel II adoption date of July 1, 2012 (t=0, which is implicit in the first differencing, covers July 1, 2010 through June 30, 2011); $D\_Aaa\text{-}Aa3_c$ ($D\_A1\text{-}Baa3\&NR_c$) [$D\_Aaa\text{-}Baa3\&NR_c$] is equal to one if the letter of credit-issuing counterparty-banks in the destination non-OECD country c are, on average, rated between Aaa and Aa3 (A1 through Baa3 or non-rated) [A1 through Baa3 or non-rated] from Moody's throughout the sample period; $D\_BASELII_t$ is equal to one for t=2, and zero otherwise; $\gamma_{c,p}$ denotes country×product (c×p) fixed effects, and $\gamma_{p,t}$ denotes product×time (p×t) fixed-effects; and ε is the error term of the OLS regression. The regressions robust standard errors are clustered at the country×time (c×t) level, and the corresponding t-statistics are provided within parentheses below coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Appendix Table A5
Summary statistics for counterparty-bank rating proxy at the country-level

	Counterparty-bank rating proxy					
	OECD Sample			Non-OECD Sample		
Ratings ranges that correspond to risk-weight categories	7.1.2011 – 6.30.2012	7.1.2012 – 6.30.2013	t-stat	7.1.2011 – 6.30.2012	7.1.2012 – 6.30.2013	t-stat
Aaa-Aa3	4.20 (0.44) [6]	4.35 (0.28) [6]	1.38	3.94 (0.34) [3]	4.14 (0.26) [3]	1.65
A1-Baa3	6.49 (1.49) [18]	7.09 (1.74) [18]	2.78 **	8.63 (1.69) [21]	8.58 (1.68) [21]	0.69
Ba1-B3				14.31 (1.39) [33]	14.30 (1.43) [33]	0.18

This Appendix table provides summary statistics for the counterparty-bank rating proxy at the country-level. Standard deviations are provided within parentheses, the number of observations within brackets. The proxy is the country-level weighted-average rating of banking institutions for which ratings from Fitch, Moody's and S&P could be found on BankScope and where the weights are bank total assets. Rating agencies' letter-scales have been turned into a numerical scale: 1 corresponds to Aaa from Moody's or AAA from Fitch or S&P, 2 to Aa1 from Moody's or AA+ from Fitch or S&P, 3 to Aa2 from Moody's or AA from Fitch or S&P, etc. For banks that have overlapping ratings from more than one of these three agencies we follow the rules imposed by the Turkish banking regulators: if a foreign counterparty bank has two agency ratings, Turkish banks have to use the worst (lower) of the two ratings, if it has three ratings Turkish banks are required to use "the better of the worst two ratings" (i.e., the middle rating). ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.