Risk-Based Capital Requirements for Banks and International Trade: Evidence from Basel II Implementation in Turkey [*]

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This version: September 10, 2014

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Abstract

We find that changes in risk-based capital requirements for banks can affect the real economy through international trade. Using a natural experiment – Basel II adoption in its Standardized Approach for all banks in Turkey as of July 1, 2012 – we investigate the impact of new risk-weights applied to commercial letters of credit (CLC) on related Turkish exports. We estimate the resulting payment-term-cost elasticity of CLC-financed exports to be between -0.5 and -1. These results are robust to the presence of zero-trade observations, bank or country credit ratings, frequency of the data (annual or quarterly), various sub-samples, and a placebo test. (100 words)

JEL: F13, F14, G21, G28

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

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[*] We would like to thank Jérôme Héricourt, Mark Carey; participants at Federal Reserve Bank of Chicago's 50th Bank Structure and Competition Conference; as well as those present during seminars at the Aalto University, Bank of France, and HEC Paris. All remaining errors are our own. Michalski and Ors have benefitted from a French National Research Agency research funding (grant number ANR-11-IDEX-0003/Labex Ecodec/ANR-11-LABX-0047).

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Effects of Basel II Implementation on Turkish Exports

1. Introduction

Many institutions involved in global trade raise serious concerns regarding the treatment of financial instruments related with exports and imports under later versions of Basel Accords 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" (March 31, 2009, p. 40). Other banking surveys indicate that (i) Basel II had a negative impact on banks' provision of trade finance for the majority of large 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, G20 stated that it would "… evaluate impact of regulatory regimes on trade finance" (G20 Seoul Summit Document, 2010). Four years later we still know very little regarding the impact of changes in capital standards on international trade due to lack of research in this area.

To the best of our knowledge, we are the first to investigate the impact of regulatory changes in Risk-Based Capital (RBC) standards on the real economy through international trade. We do so by examining changes in trade flows around the mandatory adoption of Basel II in its Standardized Approach (SA) form by all Turkish banks on July 1, 2012. More specifically, we analyze the effect of the changes in capital charges for the export-related commercial letters of credit (CLCs) held by Turkish banks on behalf of their exporter clients for shipments to 174 countries over a two-year period around the adoption date. CLCs are international trade finance instruments where the issuing bank, which is typically located in the same export-destination country, covers the foreign importer's risk of payment-default. The exporter presents the CLC for payment to its local bank, which holds it as an off-balance sheet item.[2, 3] Under Basel I and the SA version of Basel II, CLCs, as any other off-balance sheet item, are first converted into an on-balance sheet equivalent amount using a *credit*

[1] Similar fears have been raised for Basel III, which initially proposed that a 100% capital be set aside for many off-balance sheet items, including commercial 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, World Trade Organization (WTO), and 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] In some countries, such as the US, the confirmed CLCs can be sold in the secondary market as bankers' acceptances. In Turkey, there is no secondary market for export-CLCs: they are simply held as off-balance items until their maturity upon confirmation by the Turkish exporter's domestic bank.

[3] Under the Basel rules a particular CLC issued by the importer's bank at the shipment-destination country or held by the exporter's bank at the shipment-origin country generates off-balance sheet positions for both banks.

conversion factor (CCF) that multiplies the nominal value of the CLC. Then, the capital charge of the off-balance sheet CLC position is calculated by multiplying the obtained credit-equivalent amount with a *risk-weight* (RW) to adjust for counterparty exposure. In the Turkish case the CCFs applied to CLCs remained constant between 2006 and 2013, that is, both under Basel I and II.[4, 5] Our focus is on changes in exports that are due to new RWs that are applied starting July 1, 2012 with Basel II.

Under Basel I, the Turkish banks were required to apply two different RWs depending on whether the issuing banks of export-related CLCs were domiciled in an OECD or non-OECD country. In contrast, the SA version of Basel II requires that the RWs differ based on (i) the maturity of the CLCs and (ii) national regulator-defined groups of agency-rating categories following the guidelines proposed by the Bank for International Settlement (BIS). As a result, the move from Basel I to II gives us potentially four identification schemes, which are reduced to three due to the data limitations. To illustrate (more detail is provided in Section 3.1), for banks located in OECD countries for CLCs with maturities higher than three months, the associated RWs either (i) increased by 150% (from 20% to 50%) for lower investment grade (A1 to Baa3, or equivalently A+ to BBB-) rated counterparties or (ii) stayed constant at 20% for higher investment grade (Aaa to Aa3, or equivalently AAA to AA-) rated counterparties that form the base case in our difference-in-differences regressions.[6], We use Turkish export flows that are disaggregated at the industry-country of destination level.[7] For the above-mentioned group our preferred estimates indicate that Basel II induced increases in RWs led to decreases in CLC-based Turkish exports. We find comparable results with the other identification schemes.

We believe that our empirical findings are important given that there is little *direct* evidence on the effects of changes in bank capital requirements on the real economy. Most of the academic research to date has focused on the impact of capital requirements on banks' supply of loans (e.g., Berger and Udell, 1994; Brinkmann and Horvitz, 1995; and Kashyap and Stein, 2004) or their investment in financial securities (for example, Leibig et al., 2007).[8] One of our contributions is to help fill this gap by evaluating the effects of Basel II on exports, which are, for many countries, an important part of the gross domestic product (GDP).

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

[5] The CCFs remained constant even though they differ by CLC type (for example, depending on whether the exported good can be collateralized or whether CLC is confirmed; please refer to Table 1 and the discussion in Section 3.1 below for more details).

[6] In fact, for CLCs with maturities higher than three months issued by OECD banks three other categories were also possible post-Basel II: RWs (i) remained at 20% for non-rated counterparties, (ii) increased to 100% for non-investment grade and non-default rated counterparties, and (iii) went up to 150% for imminent or actual default categories (i.e., Caa1 or CCC+ and below rated counterparties). But, as explained in more detail in Section 3.1, these cases do not apply in our particular setting because of the restrictions we apply to the data.

[7] Ideally, one could use transaction-level data to investigate the changes in capital charges on CLC-based exports. Nevertheless, such detailed transaction-level data are not available to us.

[8] One exception is Brun, Fraisse, and Thesmar (2013) who find that the 2008 Basel II adoption in France increased the aggregate firm borrowing and investment allowing the preservation of jobs during a crisis period as average bank capital required for industrial loans decreased by 2%.

We also contribute to the emerging literature on the export or import payment terms and international trade flows (e.g., Antras and Foley, 2013; Auboin and Engemann, 2012; Glady and Potin, 2011; Mateut, 2012; Niepmann and Schmidt-Eisenlohr, 2013a and 2013b; Schmidt-Eisenlohr, 2013). Whereas many papers in this line of research focus on one method of trade payment (e.g., Glady and Potin, 2011; Mateut, 2012; or Niepmann and Schmidt-Eisenlohr, 2013a, 2013b), our data allow us to differentiate among trade flows given the category of payment used. In particular, we can distinguish whether the trade flows are based on CLCs described above; Cash In Advance (CIA) where the importer bears all the risk of the transaction as it pays the exporter prior to shipment; or Open Account (OA) where the exporter bears all the risk as it gets paid by the importer upon receipt of goods. As a result, our empirical set-up allows us to estimate the impact of Basel II adoption on exports that are based on these three different trade payment terms. Finally, our results also complement the findings of the recent strand of research on the impact of the Great Recession on international trade (e.g., Ahn, Amiti, and Weinstein, 2012; Asmundson et al., 2011; Levchenko, Lewis, and Tesar, 2010; Paravisini, et al., 2013) in two ways. First, we show that international trade (at least in some industries) can react strongly to changes in the implicit costs of processing CLCs. This finding is consistent with both a pricing and a rationing channel for CLC. Second, our findings suggest that changes in capital requirements that took place during the Great Recession might also have impacted CLC-based exports as banks, hence their agency ratings, weakened, increasing the trade punishment that Basel II incorporates through higher RWs.[9]

The Turkish data that we study have unique features that allow us to identify whether Basel II, in its most basic form in which it was adapted, had any impact on CLC-based exports. First, when applying the credit risk component (Pillar 1) of Basel II, the banking authorities in Turkey required that all banks under their jurisdiction only use the SA, whose constituents are public information.[10] In contrast, banking regulators in other countries typically allow their banks to choose among the three different approaches when implementing Basel II: (i) the SA, (ii) the "foundation version" of the Internal Rating Based (FIRB) approach, and (iii) the "advanced version" of the IRB (AIRB). FIRB and AIRB approaches, typically chosen by large and more sophisticated financial institutions, which are also more likely to provide international trade financing services, may differ across institutions in the capital charges that they imply for a given on- or off-balance sheet position, such as a CLC (see, for example, Financial Times, February 26, 2013). More importantly, when banks are allowed to adopt different approaches, identifying the effects of Basel II on international trade becomes more difficult, if not impossible, unless one has access to bank-and-firm level trade transaction or CLC data, which are typically proprietary and not commonly available to researchers. In contrast, the imposition

[9] For example, the 27 European Union countries adopted Basel II as of January 2008.

[10] Turkish banking authorities made it clear that the Internal Rating Based (IRB) 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.

of SA, whose components are public information, by the Turkish banking regulators to all banks under their jurisdiction provides us with a clear set of identification schemes.[11]

Second, we work with data collected by the Turkish Ministry of Customs and Trade and provided to us by the Turkish Statistics Institute. Our dataset (before we impose filters needed to conduct our analysis) covers the universe of the country's exports of manufactured goods between July 1, 2011 and June 30, 2013. Importantly, our data provide exports disaggregated by financing terms (i.e. CIA, CLC, or OA terms), country of destination, and two-digit ISIC industry level. This level of detail, which is typically unobservable in aggregate trade data, allows us to conduct our tests while controlling for unobservable fixed as well as time-varying country and industry characteristics.

Finally, Turkey is an economically relevant case to study. The country, which is a member of the OECD, WTO and the G-20, is the world's 17th largest economy, its 22nd largest exporter by value (15th largest exporter in manufactured goods that we examine) and the 14th largest importer.[12] Turkey is also 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) and its seventh largest importer.[13] 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. Even after the restrictions that we have to impose on the data to be able to properly estimate the difference-in-differences model, we still account for roughly 70% of Turkey's shipments.

To test for the potential impact of Basel II adoption on Turkish exports, we nest a difference-in-differences model into what we call a *pseudo* gravity equation. While the gravity equation is the most-widely used estimation tool in empirical international trade, various limitations of the data and the identification schemes that we use do not allow us to estimate a full-blown gravity model, even if our approach is in the same spirit. To control for time-varying country demand for foreign manufactured goods, our regressions incorporate destination-country imports for the period (excluding Turkish exports to that country) from the rest of the world. We also include distance between Turkey and the

[11] One concern regarding Basel II adoption is whether the treatment of small and medium sized enterprise (SME) counterparties under FIRB or AIRB approaches may be affecting international trade (as in ICC report, March 31, 2009, p.40 and Financial Times, February 19, 2009). In the case of Turkey, while Basel II implementation did not involve FIRB or AIRB, the banking regulator nevertheless reduced the RW from 100% to 75% for the group of SMEs with less than 50 employees and sales of less than 5 million Turkish Lira ($ 2.766 million as of July 1, 2012). The 25% decrease in the capital cost of loans to this group of Turkish SMEs is likely to increase their access to bank finance, hence their working capital, which might lead to an increase in their OA-based exports (where the firm gets paid after the shipment and thus finances of the transaction). We would expect their OA shipments to increase but their CLC-based exports not to be affected: as their loans' RW decreases to 75%, this group of SMEs would have access to potentially cheaper working capital making more funds available to finance OA transactions, but their CLC-related costs would follow the same pattern as that for larger companies. Unfortunately, we do not have a break-down of the data by firm-size.

[12] 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).

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

destination country as well as an indicator variable for adjacent countries to account for time invariant impediments to trade. We estimate alternative specifications with country-time and industry-time fixed effects to account for unobservable variables. We also use the Poisson Pseudo-Maximum Likelihood (PPML) estimator to deal with heteroskedasticity and zero-trade observations, a typical problem encountered in the empirical trade research. Inherent heteroskedasticity of trade flows to different destinations and/or omitted zero-trade flows lead to biased log-linear estimates (Santos-Silva and Tenreyro, 2006, 2012). Equally importantly, our regression models explicitly account for the three main types of trade financing terms to provide a proper counterfactual encompassing of all Turkish exports of manufactured goods. As such, we are able to factor-in the possibility that changes in domestic credit associated with Basel II might affect OA-based exports that firms finance using working capital. Consequently, we argue that the Basel II related changes in exports that are picked up by our model are driven by the supply-side of trade financing services by Turkish banks (rather than changes in the demand-side for such services by Turkish exporters). Moreover, our tests are inherently weak and potentially biased against us finding an effect as they rely on the joint hypothesis that: (i) the sovereign credit ratings or (total-assets weighted) foreign bank credit ratings are representative of the average ratings used in CLC-based exports, and (ii) Turkish banks reflect Basel II-related differential changes in capital charges to the prices of CLC clearing. These points are further clarified in Section 3.

Our findings suggest that, after Turkey's mandatory adoption of Basel II in its SA version by all of country's banks the value of CLC-settled exports to OECD countries decrease given the CLC-associated RW increase, whereas exports to non-OECD countries increase given the RW decrease. In contrast, typically we do not observe similar patterns for exports settled under other methods of payment (i.e., CIA or OA). Presuming that banks reflected the changes in CLC-confirmation costs to their pricing, at least to a certain extent, we can calculate the elasticity of CLC-intermediated exports to the changes in the bank's cost of capital to be between -0.5 and -1.0. These findings suggest that the reaction of trade values to changes in the cost of trade finance are economically relevant. Given that in the pre-Basel II period the CLC-export shares are on average 6.4% for the OECD countries and 17.9% for the non-OECD countries, using a back-of-the-envelope calculation and presuming that there are no other changes to shipments, we estimate that the overall elasticity of exports to RW changes was between -0.032 and -0.179. In other words, a 1% increase in trade costs associated with RW changes lead to 0.03% to 0.18% decrease in trade flows. These estimates are comparable to those found Paravisini et al. (2013) for the reaction of Peru's total trade to financial shocks during the Great Recession. We also find evidence that is consistent with a rationing story. CLC-based exports to countries with speculative-grade (below Baa3 and BBB-) ratings decrease. This finding is consistent with Turkish banks' imposing internal credit exposure limits for non-investment grade counterparties after Basel II. These results are robust to the use of equally-weighted average bank ratings instead of TA-weighted bank ratings as a proxy for average ratings involved in Turkish exports to different

destinations, the presence of zero-trade observations, frequency of the data (annual or quarterly), various sub-samples, and a placebo test.

Our paper proceeds as follows. Section 2 provides a survey of the academic research that is relevant for our work. In Section 3 (i) we detail our identification scheme, (ii) introduce the empirical specifications that we use in our analysis, and (iii) provide information and summary statistics on our data. Section 4 presents our main empirical results, which are followed by robustness checks in Section 5. Section 6 concludes the paper.

2. Literature Review

Our paper draws upon, and contributes to, three strands of research. The first of these examines the impact of Basel Accord capital requirements on banks' lending behavior and the real economy. Most of the papers in this area of research examine the impact of capital requirements on banks' loan provision. Early papers by Peek and Rosengren (1995a and 1995b) show that New England banks that are subjected to a "capital crunch" (shortage of capital under higher capital ratios, which need not equal RBC requirements) decrease their lending more. However, Berger and Udell (1994) use a larger panel dataset, control for alternative explanations, and find little evidence that the RBC requirements that U.S. regulators imposed during the early 1990s explain the credit crunch that followed. More recently, Kashyap and Stein (2004) show that simulated capital charges, and hence the lending costs, increase under Basel II compared to Basel I. Berger (2006) examines the potential impact of AIRB adoptions on SME loans, and concludes that the economic effect is likely to be unimportant as small banking organizations are unlikely to adopt AIRB and the larger institutions that do so are much less likely to make typical SME loans. On the investments side, Liebig et al. (2007) find that German banks' sovereign lending to emerging economies is little affected by RBC requirements after Basel II adoption. All of the above papers focus on the supply of loans by banks but do not directly test for the effects of capital requirements on the real sector.

One exception is Brun, Fraisse and Thesmar (2013) who use matched bank-and-firm loan-level French data to assess the impact of Basel II adoption in 2008 on bank lending as well as on corporate activity. Using regulator-collected data covering multiple bank-firm relationships, these authors are able to control for unobservable bank heterogeneity as well as unobservable changes in firm credit demand. Moreover, having access to supervisor collected proprietary data, Brun, Fraisse and Thesmar (2013) are able to account for banks' adoption of SA, IRB or AIRB approaches and their internal credit risk assessments. They find that following the 2008 Basel II adoption the capital requirements on industrial decreased by 2% leading to a 10% increase in the average loan size. As a result, the borrowing by French firms went up by roughly 12 billion euros, a 1.5% increase. Similarly, investment increased by 0.5%, leading to the preservation 235,000 jobs, 1% of the aggregate French employment, during the crisis period. In contrast, our findings provide additional evidence that Basel II adoption also affects the real sector through another channel, i.e., via its impact on trade flows.

Second, our work is also related with the recent and important research that examines the role of financial intermediaries in international trade. One strand of this area of research examines the role of banking integration on trade. Michalski and Ors (2012) find that trade flows between US states grow as financial integration across-state borders increases following deregulations of interstate banking entry restrictions. In follow up work, Michalski and Ors (2013) find that entry by domestic banks with international assets as a US state becomes more financially integrated with the rest of that country leads to higher state-level exports to foreign destinations. In a related paper Hale et al. (2013) find that international trade increases as new international banking links, as proxied by international syndicated loans, are established. Another series of papers examine the impact of shocks to banks on international trade. For example, Ronci (2004) shows that a fall in trade financing following a domestic banking crisis leads to lower exports. Amiti and Weinstein (2011) find that one third of the drop in the trade-to-GDP ratio for Japan in the 1990s can be explained by the poor financial health of the main banks of large Japanese exporters. Focusing on exports by small US firms during the latest downturn, Peek (2013) finds that export share of SMEs decreases with deteriorating bank financial health. Using export transactions data, Paravisini et al. (2013) show that the negative credit supply shocks experienced by Peruvian banks during the 2008 financial crisis account for 15% of the drop in the country's exports in the same period. Ahn (2013) conducts a similar exercise for Colombia using CLC-financed imports. Another strand of literature in this larger area examines the impact of credit constraints on exports. Chaney (2013) provides a model with liquidity-constrained exporters. Relying on a survey of Italian firms, Minetti and Zhu (2011) use the differences in historical Italian banking regulations as an instrument and find that firm-level exports are negatively affected by credit constraints. According to Chor and Manova (2012) external-finance dependent sectors in countries with adverse credit conditions experienced larger falls in their exports to the US during the 2008 crisis.[14] We complement this area of literature by examining the effects of a regulatory change that affects the capital charge (the implicit cost) of holding CLCs by Turkish banks, which in turn might affect exporters' behavior when serving different destination countries.

Finally, our paper also contributes to a recently emerging area of research that examines the role of different methods of payment between firms in international trade, on which little is known, in contrast to the larger literature on (domestic) trade-finance (e.g., Giannetti et al. 2011 or Klapper et al. 2012). A number of papers in this line of research focus on one type of international-trade payment at a time. Mateut (2012) finds that CIA payments, which can be linked with both firm and industry characteristics, can be used to reduce default risk in international trade. Glady and Potin (2011) develop a model and find empirical evidence for the fact that asymmetric information and difficulties in contract enforcement increase CLC default risk, which financial intermediaries are able to reduce.

[14] That said, Levchenko et al. (2010) and Eaton et al. (2011) suggest that the drop in trade finance played a minor role during the trade collapse of 2008–2009, with the steep decline in trade-to-GDP ratio being linked with the lack of demand for intermediate or durable goods during the crisis.

Auboin and Engemann (2012) find that 1% increase in international trade credit for a given country leads to a 0.4% increase in its real imports. Other papers try to characterize trade-offs that might be involved across different international trade payment terms. Demir and Javorcik (2014) use a similar Turkish exports dataset to ours and find that exporter financed OA-based exports increase, relative to CIA- or CLC-based, with institutional quality, banking efficiency, and level of market competition in the importing country. They also find stronger effects for differentiated products. Schmidt-Eisenlohr (2013) formulates a model that links trade financing terms to the financial market characteristics and the contracting environments of the countries in which the exporter and the importer are located. His model's predictions are supported in aggregated trade data. Antras and Foley (2013) characterize export transactions data from a large US poultry firm and rationalize the empirical patterns that they observe in an extension of the model developed by Schmidt-Eisenlohr (2013). While we cannot directly examine trade finance choices, as Antras and Foley (2013) do, due to the aggregate (at the country-industry) level of the data, we nevertheless allow for changes (around Basel II adoption) in CIA, CLC or OA-financed trade flows to the same risk-weight category destinations. As a result (i) we can account for changes in trade flows in a comprehensive way, and (ii) we can provide tests as to whether changes in the (implicit) cost of export-related CLCs affect shipments that are based on other payment terms. In the next section, we detail our identification scheme, the empirical specifications that we use, and the data that are at our disposal.

3. Identification, empirical specifications, and the data

3.1. Identification

As indicated in the Introduction, our three identification schemes rely on the Basel II induced changes in export-CLC capital charges, which are equal to the letter's nominal value times CCF times RW. In the Turkish case, the same CCFs apply both under Basel I and the SA version of Basel II (Articles 5 of BDDK Directives of November 1, 2006 and June 28, 2012). As detailed in Table 1, the CCF can be equal to 100% (for confirmed, irrevocable export-CLCs), 20% (for irrevocable CLCs with less than one year of maturity and for which the exported good serves as collateral), 0% (for non-binding CLCs that do not require a payment), and 50% (for any irrevocable CLC that does not fall in the previous three categories). Unfortunately, the Ministry of Customs and Trade does not collect this information in the international trade transaction forms, which are the basis of the aggregate data made available to us. However, this is not a major concern for a number of reasons. First, CLCs are typically issued by counterparty banks that are located in the same country as the foreign importer. Given that our final dataset includes 174 countries, it is unlikely that importers in these destinations would change the type of the CLCs that they request from their bank in a coordinated way so as to attenuate capital charges (hence higher CLC clearing fees) faced by Turkish exporters in cases when the RWs increased with Basel II. Second, even if such an implausible scenario were to occur, the effect would go against us finding any effect, as it would decrease (counterbalance) the increases in

Turkish banks' capital charges for export-CLCs. Third, we look at a relatively short period of time (two years centered around July 1, 2012) during which CLC-type composition is unlikely to vary by large proportions at the country or industry level. As a result, our country and industry fixed effects would soak up unobservable CLC-type composition effects that remain constant over time. Fourth, even if CLC-type composition did vary over time, we employ country-time as well as industry-time fixed effects that would account for unobservable variations in exports that are due to changes in CLC-type composition (albeit separately at the country and industry levels, as our data are at the country-industry-year or quarter level). As a result, our focus in on Basel II-related changes in RWs.

Table 2 provides a summary of the RWs before and after Basel II adoption. Under Basel I, RWs applied by Turkish banks to export-CLCs were either equal to 20% for OECD-country based commercial banks or to 100% for non-OECD-country based commercial banks (BDDK Directive, November 1, 2006, Supplement 1, Article c) and, Supplement 2, Article VII). With Basel II the RWs for export-CLCs vary depending on (i) whether the remaining maturity of the letter is longer or shorter than three months and (ii) the groups of counterparty bank agency-rating categories.[15] Unfortunately, the customs forms, based on which the aggregated data are collected by the Ministry of Customs and Trade do not collect counterparty foreign bank information. Instead, we rely on TA-weighted average commercial bank ratings for each country.[16] A potential downside of average commercial bank ratings is that we would not be accounting for CLCs issued by non-rated foreign banks. But we have to live with this shortcoming, as we have no way of accounting for all financial institutions in a given country.

Given our proxy for foreign counterparty banks' ratings, our identification strategies may be better understood by going through two simple examples that are based on Table 1. Our hypothesis is that the banks would at least partially reflect these changes into CLC-clearing prices and/or ration

[15] In fact, Basel II provided domestic banking authorities with two "options" for the RWs for (on and off balance sheet) foreign bank liabilities. The Turkish banking regulator, opted for the "option 2" (BDDK, July 2007, p. 10), which takes into account counterparty commercial banks' agency ratings and whether remaining maturity is longer or shorter than three months, as described above. The "option 1" suggests that "…all [commercial] banks incorporated in a given country will be assigned a risk weight one category less favourable than that assigned to claims on the sovereign of that country. However, for claims on banks in countries with sovereigns rated BB+ to B- and on banks in unrated countries the risk weight will be capped at 100%." (BIS, June 2004, Article 61, page 29). The corresponding benchmark-RWs, *irrespective* of maturity of the off-balance sheet position are as follows: 0.20 for AAA and AA-, 0.50 for A+ to A-, 1.00 for BBB+ through B- and for non-rated sovereigns, and 1.50 for CCC and below (BIS, June 2004, Article 63, page 30).

[16] An alternative to TA-weighted bank ratings is to use equally-weighted bank ratings per country so as not to penalize smaller financial institutions that might specialize in foreign-trade financing. We find similar results when we do so. Another alternative is to use sovereign long-term credit ratings (as we did in an earlier version of this paper). However, it is not clear whether sovereign ratings would provide as good a proxy as the average of CLC-issuing banks' ratings in that country. Although the Turkish banking regulator initially chose *not* to adopt the 2011 BIS recommendation to remove the so-called "sovereign floor" for financial instruments (such as CLCs) used in international trade (Basel Committee on Banking Supervision, October 2011; BDDK Directive of November 1, 2006 Article 4.(6); and BDDK Directive of June 28, 2012, Supplement 1, Part I, Section 6, Article 26 under sub-section 6.2), during the implementation it allowed banks to apply RWs that are lower than that of the sovereign where the bank is domiciled if the institutional ratings were to be higher. We nevertheless provide results based on sovereign ratings in Appendix Table A2.

(increase) CLC-confirmations, and exporters would react accordingly: CLC-based exports would increase (decrease) in the cases when capital charges decrease (increase).

First, suppose that a Turkish bank gets a request from a domestic customer to clear an export-related confirmed-CLC of $ 1 million (approximately equal to 1.8 million Turkish Liras [TL] on July 2, 2012) issued by the importer's bank and with a remaining maturity of more than three months. Prior to July 1, 2012 under Basel I, if the counterparty bank was located in an OECD country, holding this export-related CLC would have required that the Turkish bank set aside $ 24,000 (approximately TL 43,400) in capital, *irrespective* of the risk of the counterparty-bank issuing the CLC.[17] After July 1, 2012, under Basel II, the same CLC's capital charge would depend on the OECD-based counterparty bank's agency rating (please refer to columns A and D of Table 1). Suppose first that the agency rating for (the long-term foreign currency denominated debt of) the country of domiciliation of the counterparty bank's is in the Aaa to Aa3 range (i.e., among top four investment grade categories) or non-rated. Then the capital charge would remain equal to $ 24,000 as the associated RW would remain equal to 0.20 after Basel II adoption for what is our control group. Now suppose that the sovereign rating is between A1 and Baa3 (i.e., among the lower six investment grade categories). The corresponding RW would increase to 50%; hence the capital charge would go up to $ 60,000, a 150% increase.[18] In such cases, we would expect the CLC-clearing cost to increase, potentially affecting CLC-based exports to countries whose banks are rated A1 and Baa3 on average, compared to the control group. It could also be that Turkish banks simply ration more often CLC-requests for A1 and Baa3 rated counterparties.

We could also do a similar exercise when the same export-related confirmed-CLC is issued by a counterparty bank domiciled in a non-OECD country (please refer to columns B and D of Table 1). Under Basel I, the associated capital that needs to be set aside by the Turkish bank was equal to $ 120,000. [19] Under Basel II the capital charge would decline by 80% to $ 24,000 if the counterparty rating is Aaa to Aa3; by 50% to $ 60,000 if the counterparty rating is between A1 and Baa3 or if the counterparty is non-rated; would not change if the counterparty rating is between Ba1 and B3 (which forms the control group in this case); and would *increase* by 50% to $ 180,000 if the counterparty rating is Caa1 and below. Finally, we have a third identification scheme for CLCs to non-OECD

[17] $ 24,000 = $ 1,000,000×1.00×0.20×0.12, where CCF is equal to 100% (for a confirmed export-CLC), RW is equal to 20% (for an OECD counterparty under Basel I), and the minimum Tier 1+Tier 2 capital ratio is equal to 12% (as required by Turkish banking regulators).

[18] For OECD countries, the other two possibilities (where the RW increases to 100% (below investment grade but non-default ratings) or to 150% (imminent or actual default) are not covered here as such cases are eliminated from our sample given the data restrictions we have to impose for a proper difference-in-differences estimation.

[19] $ 120,000 = $ 1,000,000×1.00×1.00×0.12, where CCF is equal to 100% (for a confirmed export-CLC), RW is equal to 100% (for an 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).

countries with remaining maturities of less than three months (as depicted by the combination of columns B and C in Table 2).[20]

While we have three clear identification schemes (two of which applies to one dataset, the third to a completely different dataset) that are well specified at a given point in time (i.e., Basel II adoption date of July 1, 2012) across export-destinations (due to differently rated destination country groups being differently affected depending on whether they are OECD or non-OECD member countries) and depending on the average remaining maturities of CLCs, changes to risk-weights need not necessarily affect trade transactions. First, Turkish banks need not fully reflect Basel II related changes to capital charges into their prices for CLC-clearing. Although some of bankers that we spoke to indicated that the prices of CLC-related services were affected by Basel II adoption, others stated that large firms customers with repeat business were less likely to be affected than SMEs with less frequent export transactions. As a result, RW changes that we use in our identifications may have had only a partial impact, if any, on related export flows.

Second, Turkish banks, unlike their EU or US counterparts during the same period, were well capitalized as of June 2012. Their 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.[21] If Turkish banks internalized the capital cost 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 any changes in the related trade flows.

Third, to reiterate a point made earlier, the aggregated country-industry exports flows that we use do not allow us to trace the risk of the bank counterparty. Instead we use TA-weighted long-term (foreign currency denominated) debt ratings (as mandated by the SA version of Basel II) for all banks in a given country as a proxy for the average counterparty foreign-bank risk rating. As a result, due to data restrictions our tests are not as precise as we would like them to be.

Fourth, we do not know the average export-CLCs' maturities for Turkish exports. This further weakens our tests because under Basel II's SA, the risk-weights (hence capital requirements) per rating category differ depending on whether the CLC's *remaining* maturity is longer or shorter than

[20] There are potentially two additional identification schemes that we cannot use. First, as the combination of columns A and C of Table 2 indicates, there is a fourth potential weighting scheme for CLCs that have less than three months of remaining maturity and issued by banks domiciled in OECD countries. But, after the filters that we apply to our data for proper difference-in-differences estimation, we have no export-destination country left in the non-investment grade category for which the RW changes when moving from column A to column C in Table 2. Second, Basel II Directive of June 28, 2012 suggests a different ideantification scheme based on original short-term agency ratings. The directive proposes a different set of RWs for receivables that have less than three months of remaining maturity and for which one or more agencies have issued a receivable-specific (rather than bank-specific) original short-term rating (BDDK Directive of June 28, 2012, Supplement 1, Part I, Section 6.4, Articles 33 and 34 as well Section 14., Article 64). But this part of the Directive is non-applicable in our case because banks do not request an (short- or long-term) agency rating for the CLCs that they issue. As a result, we are left with the three identification schemes based on long-term bank ratings as described above.

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

three months. Existing surveys on the issue provide mixed results. For example, a 2009 SWIFT report based on a different sample suggests that roughly 50% of CLCs mature in 60 days and that almost 90% expire in less than 90 days (SWIFT, October 2009). In contrast, a 2011 ICC report indicates that, based on a dataset of more than 11.4 million transactions over 2005-2010, the unweighted average "life-cycle" (i.e., maturity) of confirmed export-CLCs (including both CLCs requiring on-sight as well as deferred payments) was 103 days (International Chamber of Commerce, 26 October, 2011, p. 16). These surveys' results need not necessarily apply to CLCs used in Turkish exports. So in our regressions we presume that majority of CLCs have maturities are either longer or shorter than three months, and run separate regressions accordingly. Of course, in reality the results to be obtained if actual maturity structures were to be known would be a combination of these two sets of results. As a result when interpreting our results with the specifications described in the next section, one should keep in mind that our tests are inherently weak and tilted against us finding any effect for the reasons described above.

Moreover, we remain cautious in the interpretation of our results for another reason as well. For example, we cannot make direct welfare calculations for the totality of Turkish exports. This is because we are restricted in the inferences we can draw from the industry-country level data given (i) the restrictions we need to impose on them for a proper difference-in-differences estimation, and (ii) the identification schemes that differ for OECD and non-OECD countries.

3.2. Empirical specifications

3.2.1. Log-linear models

To conduct our analysis we estimate a pseudo gravity model in which we embed a difference-in-differences model. A typical gravity equation relates international trade flows of countries with a set of predictors commonly used in the empirical research on international trade. In our case, we estimate what we call a pseudo gravity model, because (i) we are only interested in Turkish exports to different destinations (as opposed to different countries' exports to each other), (ii) we cannot account for the totality of the Turkish exports (for reasons mentioned above and further clarified below), and (iii) some of the variables (such as import demand at the destination country excluding Turkish exports) that we use are not standard gravity equation explanatory variables (such as GDP).[22] At any rate, our focus is on the difference-in-differences part of the empirical model that allows us to test whether Basel II induced risk-weight changes across different groups of rating classes had any impact on exports to the corresponding groups of destination countries.

To describe our empirical strategy we start with a log-linear model. Even though log-linear gravity models are known to suffer from a series of weaknesses (detailed in Section 3.2.2 below), we

[22] GDP data, which are the standard measure of destination import-demand, are only available on an annual basis for most countries. Given that our natural experiment occurs in the middle of the year we had to resort to a different variable than the GDP so that we can aggregate quarterly data into annual periods accordingly,.

nevertheless rely on them here (before moving on to our preferred models) as they provide a simple benchmark that is easy to describe. Our starting point is the following log-linear empirical pseudo gravity equation:

$$ln(EXPORTS_{c,i,t}) = \alpha_0 + \alpha_1\, ln(IMPORTS_EX_TUR_{c,t}) + \alpha_2\, ln(DISTANCE_c) + \alpha_3\, D_ADJACENT_c + \delta_i + \varepsilon_{c,i,t} \quad (1)$$

where, subscript c denotes destination country, subscript i denotes two-digit industry segment, subscript t denotes time period, and prefix $D_$ denotes indicator variables; ln$(EXPORTS_{c,i,t})$, the dependent variable, is the natural logarithm of Turkish *CLC-based* exports (in US dollars) to country c in industry sector i during period t; $ln(IMPORTS_EX_TUR_{c,t})$, which controls for the destination-country import-demand, is the natural logarithm of total imports to country c in during period t after excluding of Turkish exports to that country; $ln(DISTANCE_c)$ is the geographical distance between Turkey and country c; $D_ADJACENT_c$ is an indicator variable that is equal to one if country c has a land-border with Turkey, and zero otherwise; δ_i is an industry fixed effect; and $\varepsilon_{c,i,t}$ is the regression error term. In Eq. (1) coefficients α_1 and α_2 are elasticities that correspond to continuous variables $ln(IMPORTS_EX_TUR_{c,t})$ and $ln(DISTANCE_c)$. In contrast, the interpretation of the coefficient α_3 for the indicator variable $D_ADJACENT_c$ requires that we calculate the related incidence ratio (i.e., exp(α_3) – 1).

Next, we embed a difference-in-differences model into Eq. (1). We first focus on the OECD sample, based on the assumption that all CLCs used in have remaining maturities longer than three months, for which we obtain:

$$ln(EXPORTS_{c,i,t}) = \alpha_0 + \alpha_1\, ln(IMPORTS_EX_TUR_{c,t}) + \alpha_2\, ln(DISTANCE_c) + \alpha_3\, D_ADJACENT_c$$
$$+ \beta_1\, D_A1\text{-}Baa3_c + \beta_2\, D_BASELII_t + \beta_3\, D_A1\text{-}Baa3_c \times D_BASELII_t + \delta_i + \varepsilon_{c,i,t} \quad (2)$$

where, $D_A1\text{-}Baa3_c$ is equal to one if 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) for which the RW increases from 20% to 50%, and zero otherwise;[23] $D_BASELII_t$ is equal to one for the period(s) following Basel II adoption on July 1, 2012, and zero otherwise; with the remaining variables being as described previously. In order to be able to estimate proper difference-in-differences models (in which the rating agency, i.e., RW, category-level unobservables are captured by the same set of constants throughout the estimation period) we require that export-destination countries in our sample belong to the same rating range (as defined by the SA version of Basel II, i.e., per given row of Table 1) between July 1, 2011 and June 30, 2013. As a result, the indicator variable that correspond to rating-class (i.e., RW) category has only a country subscript but

[23] While the 50% RW also applies to non-rated countries, there is no OECD-member state whose long-term sovereign debt in foreign currency that has not been rated by an agency.

no time subscript. The omitted (i.e., the base case) category of bank ratings is Aaa to Aa3 according to Moody's (AAA through AA- according to S&P or Fitch) for which the risk-weight is at 20% under *both* Basel I and II for OECD countries. It should be noted that Ba1 to B3 (BB+ to B-), Caa1 (CCC+) and below rating categories as well as non-rated group are excluded from the model because no OECD country remains in one of these ranges throughout the sample period. For example, Greece is excluded from the OECD sample because during our sample period the proxies for its banks' ratings (sovereign rating as well as TA-weighted average bank rating) move up from default range (for which the RW is 150% if we consider CLCs with maturities higher than three months) to speculative grade range (for which the Basel II RW for maturities higher than three months is 100%). There are no OECD-member countries or country banks that are not rated during the period we study.

The coefficient estimates of interest are β_1 through β_3. Since these coefficients correspond to indicator variables or their interactions, our interpretation of their impact requires calculation of incidence ratios as described above. Coefficient estimate β_1 measures pre-Basel II difference, if any, in CLC-based exports for the group of ("treated") OECD countries for which RW has eventually increased from 20% to 50%. Coefficient estimate β_2 measures the change in CLC-based exports post-Basel II for the base-case ("non-treated") OECD countries for which the RW remains constant at 20% throughout the period we examine. The coefficient β_3 is an estimate of the post-Basel II change in CLC-based exports for the "treated" OECD countries for which CLC related capital charges became 150% more expensive in terms of capital charges with respect to the "non-treated" group of OECD countries for which the risk-weight remains at 20% after adoption. Our hypothesis suggests that β_3 should be negative: as the capital charge for CLCs increases from 20% to 50%, exports to A1 to Baa3 rated OECD countries are expected to decrease in the post-Basel II period.

Of course, besides CLC, trade financing payment terms also include OA (where the exporter gets paid upon receipt of goods and bears the transaction's risk) and CIA (where the importer pays in advance and bears the transaction's risk). Since CLC-based exports correspond to roughly one-tenth of Turkey's exports, by excluding OA and CIA transactions we would not be taking into account the proper counterfactuals. Put differently, Eq. (2) estimates may be biased, hence our economic inferences might be wrong, because we leave out almost 90% of Turkish shipments to countries that are in our sample. As a result, we estimate a triple-differences model after modifying Eq. (2) as follows:

$$\begin{aligned} ln(EXPORTS_{c,i,t}) = {} & \alpha_0 + \alpha_1\, ln(IMPORTS_EX_TUR_{c,t}) + \alpha_2\, ln(DISTANCE_c) + \alpha_3\, D_ADJACENT_c \\ & + \gamma_1\, D_A1\text{-}Baa3_c + \gamma_2\, D_BASELII_t + \gamma_3\, D_A1\text{-}Baa3_c \times D_BASELII_t \\ & + [\,\beta_1\, D_A1\text{-}Baa3_c + \beta_2\, D_BASELII_t + \beta_3\, D_A1\text{-}Baa3_c \times D_BASELII_t\,] \times D_CLC_{c,i,t} \\ & + [\lambda_1\, D_A1\text{-}Baa3_c + \lambda_2\, D_BASELII_t + \lambda_3\, D_A1\text{-}Baa3_c \times D_BASELII_t\,] \times D_CIA_{c,i,t} + \delta_i + \varepsilon_{c,i,t} \quad (3) \end{aligned}$$

where, $ln(EXPORTS_{c,i,t})$ is now the natural logarithm of the dollar value of exports, which can be financed through OA, CLC or CIA, from Turkey to an OECD destination country *c* in industry i for period *t*; with all the other variables defined as above, but with the addition of $D_CLC_{c,,i,t}$ ($D_CIA_{c,i,t,}$) which is an indicator variable that is equal to one for exports financed through CLC (CIA), and zero otherwise. In this specification the OA transactions form the base case (as this is the most often used method of payment in Turkish exports, see for example Demir, 2014). It should be noted that, despite the addition of OA- and CIA-based exports, the interpretations of the coefficient estimates of interest, namely β_1 through β_3, remain the same as in the case of Eq. (2) where we considered only CLC-based exports. A priori, we do not expect Basel II to affect OA and CIA financed Turkish exports. In the next section we describe the weaknesses that plague log-linear gravity equation models and describe the alternative Poisson Pseudo-Maximum Likelihood (PPML) estimator that is used in the empirical trade research.

3.2.2. Poisson Pseudo Maximum Likelihood regression models

As it is the case with most trade datasets, our data exhibit heteroskedasticity and contain many zero export transactions. First of these problems arises because the size of trade flows vary by country (as these differ in GDP and distance to Turkey that affects the demand for Turkish goods) and industry segment. The second problem is due to the fact that Turkey does not export in all periods in all industries to all the countries with which it trades using all three types of trade finance methods (i.e., OA, CIA, and CLC). Santos-Silva and Tenreyro (2006) show (in a cross-section) that both of these problems lead to biased and potentially inconsistent log-linear gravity model estimates when OLS is used.[24] Instead, Santos-Silva and Tenreyro (2006, 2011) suggest using Poisson or PPML regressions, which are becoming the norm in empirical trade. The Poisson regression assumes that the data are not over-dispersed, i.e., that the ratio of the mean of the data to its standard deviation is close to one. We observe that this is not the case for our data: the industry-country level Turkish exports data are highly dispersed. In the latter case PPML provides a more flexible approach than Poisson regression by allowing the variance to be proportional to the mean of the data.[25] As a result, to accommodate zero-exports and to obtain unbiased estimates for our gravity equation, we estimate the following PPML version of Eq. (3) for the OECD countries:

$$EXPORTS_{c,i,t} = exp\ \{\ \alpha_0 + \alpha_1\ ln(IMPORTS_EX_TUR_{c,t}) + \alpha_2\ ln(DISTANCE_c) + \alpha_3\ D_ADJACENT_c + \gamma_1\ D_A1\text{-}Baa3_c + \gamma_2\ D_BASELII_t + \gamma_3\ D_A1\text{-}Baa3_c \times D_BASELII_t$$

[24] One solution is to transform the dependent variable by adding $ 1 to all export flows and then take their natural logarithm, i.e., $\ln(1+EXPORTS_{c,t})$. However, Santos-Silva and Tenreyro (2006, 2011) also show that a gravity regression in which the dependent variable is transformed by adding one leads to even higher biases in the estimates.

[25] A third alternative is the negative binomial regression, which we do not use because it is unit sensitive (i.e., coefficient estimates differ depending whether dollars, thousands of dollars or millions of dollars are used).

$$+[\beta_1 D_A1\text{-}Baa3_c + \beta_2 D_BASELII_t + \beta_3 D_A1\text{-}Baa3_c \times D_BASELII_t] \times D_CLC_{c,i,t}$$
$$+[\lambda_1 D_A1\text{-}Baa3_c + \lambda_2 D_BASELII_t + \lambda_3 D_A1\text{-}Baa3_c \times D_BASELII_t] \times D_CIA_{c,i,t}\} + \delta_i + \varepsilon_{c,i,t} \quad (4)$$

where *exp{.}* denotes the exponential function and the rest of the variables are defined as above. It should be noted that the interpretations of the PPML coefficient estimates are similar to their log-linear counterparts.

For the non-OECD countries (for which the rating, i.e., RW, groups differ), still assuming that CLCs have longer than three months of maturity on average, we estimate the following version of Eq. (4):

$$EXPORTS_{c,i,t} = exp\{ \alpha_0 + \alpha_1 \ln(IMPORTS_EX_TUR_{c,t}) + \alpha_2 \ln(DISTANCE_c) + \alpha_3 D_ADJACENT_c$$
$$+\gamma_1 D_Aaa\text{-}Aa3_c + \gamma_2 D_A1\text{-}Baa3\&NR_c + \gamma_3 D_BASELII_t + (\gamma_4 D_Aaa\text{-}Aa3_c + \gamma_5 D_A1\text{-}Baa3\&NR_c) \times D_BASELII_t$$
$$+[\beta_1 D_Aaa\text{-}Aa3_c + \beta_2 D_A1\text{-}Baa3\&NR_c + \beta_3 D_BASELII_t + (\beta_4 D_Aaa\text{-}Aa3_c + \beta_5 D_A1\text{-}Baa3\&NR_c) \times D_BASELII_t] \times D_LC_{c,i,t}$$
$$+[\lambda_1 D_Aaa\text{-}Aa3_c + \lambda_2 D_A1\text{-}Baa3\&NR_c + \lambda_3 D_BASELII_t + (\lambda_4 D_Aaa\text{-}Aa3_c + \lambda_5 D_A1\text{-}Baa3\&NR_c) \times D_BASELII_t] \times D_CIA_{c,i,t}\}$$
$$+ \delta_i + \varepsilon_{c,i,t} \quad (5)$$

where, $D_Aaa\text{-}Aa3_c$ is equal to one if banks in the destination non-OECD country *c* have, on average, ratings that are between Aaa to Aa3 (for which RW 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 non-OECD country *c*'s banks have a rating between A1 and Baa3 on average or they are not rated by any of the three rating agencies between July 2011 and June 2013 (for which groups the RW drops from 100% to 50%), and zero otherwise; with the remaining variables being as described above. For non-OECD countries, the expected signs of the coefficient estimates of interest for the triple interaction are now *positive*: as RW applied to a CLC from a Aaa to Aa3 (A1 to Baa3) rated counterparty decreases from 100% to 20% (50%), our hypothesis suggests that related exports would increase with respect to the base-case category (Ba1 to B3 rated counterparties) for which RW remains constant at 100%. In Eq. (5) Caa1 (CCC+) and below rated countries corresponding to (impending or realized) default are excluded from the model because there is only one country (Cuba) in that category.

To account for the possibility that export-CLCs can have, on average, maturities less than three months we estimate the following version of Eq. (5) for non-OECD countries:[26]

$$EXPORTS_{c,i,t} = exp\{ \alpha_0 + \alpha_1 \ln(IMPORTS_EX_TUR_{c,t}) + \alpha_2 \ln(DISTANCE_c) + \alpha_3 D_ADJACENT_c$$
$$+\gamma_1 D_Aaa\text{-}Baa3\&NR_c + \gamma_2 D_BASELII_t + \gamma_3 D_Aaa\text{-}Baa3\&NR_c \times D_BASELII_t$$
$$+[\beta_1 D_Aaa\text{-}Baa3\&NR_c + \beta_2 D_BASELII_t + \beta_3 D_Aaa\text{-}Baa3\&NR_c \times D_BASELII_t] \times D_LC_{c,i,t}$$
$$+[\lambda_1 D_Aaa\text{-}Baa3\&NR_c + \lambda_2 D_BASELII_t + \lambda_3 D_Aaa\text{-}Baa3\&NR_c \times D_BASELII_t] \times D_CIA_{c,i,t}\} + \delta_i + \varepsilon_{c,i,t} \quad (6)$$

[26] Given the data restrictions that we impose, there is no OECD country for which we can run a similar regression since the RW remains constant at 0.20 (for countries with available data) under both Basel I and II.

Finally, in other versions of equations (3), (4), (5) and (6) we introduce either (i) country fixed effects (together with industry fixed effects), (ii) country-time fixed effects (alongside with industry fixed effects), and (iii) country-time as well industry-time fixed effects. In the first case the variables with only a country subscript *c* (i.e., $ln(DISTANCE_c)$, $D_ADJACENT_c$ and $D_Aaa\text{-}Baa3\&NR_c$) drop out. In the second and third cases, $ln(IMPORTS_EX_TUR_{c,t})$, $D_BASELII_t$ and $D_Aaa\text{-}Baa3\&NR_c \times D_BASELII_t$ drop out as well. In the next section we describe the data with which we estimate equations (3) through (6).

3.3. Data

Our dataset is constructed from four different sources. The Turkish Statistical Institute (TSI) provided the exports data. These are part of the all-comprehensive (and confidential) international trade dataset that is maintained by the Turkish Ministry of Customs and Trade based on individual shipment documents that are filed electronically. We obtained quarterly exports data between July 1, 2011 and June 30, 2013 aggregated by country of destination, two-digit International Standard Industrial Classification (ISIC) category and by trade financing type (CIA, CLC and OA). We further aggregate these quarterly data to come up with one year of pre- and one year of post-Basel II adoption exports data at the industry-country level. In other words, we use quarterly data so we can aggregate 2011Q3-2012Q2 and 2012Q3-2013Q2 country-industry data into two annual pre- and post-Basel II periods, respectively. We do this aggregation in order to control for (i) seasonality effects that might otherwise be picked by the difference-in-differences model's time interactions and (ii) potential serial correlation in the error terms in the panel (Bertrand, Duflo, Mullanaithan, 2004), but also (iii) to attenuate the problem of zero-trade observations. We restrict ourselves to shipments by the manufacturing sectors, which formed 93.92% of Turkish *goods* exported in 2012. These exclude barter transactions and goods that are re-exported from special trade zones established within Turkish borders. We impose the following filters on the exports data. First, as described above, we exclude countries for which sovereign debt ratings changed in such a way that the countries in question moved from one risk-weight category into another (say, from AAA-Aa3 into A1-A3) some time between July 1, 2011 and June 30, 2013.[27] Second, we exclude Cuba, which retains is Caa1 long-term foreign currency sovereign debt rating from Moody's during our sample period, because it is (i) the only country in the highest (150%) risk-weight category and (ii) a marginal destination for Turkish exports (hence with many zero-observations). Third, we also exclude Iran and Syria from our dataset. We drop Iran because, during the period of our study, Iran was subjected to an international embargo,

[27] The following countries are excluded for this reason: Azerbaijan, Barbados, Belize, Benin, Chile, Croatia, Cyprus, Egypt, Greece, Grenada, Hungary, Indonesia, Ireland, Italy, Jamaica, Korea, Pakistan, Philippines, Portugal, Romania, Slovenia, Spain, Tunisia, and Uruguay. Exports to these destinations correspond to 23% of the total value of exports in the original data.

which tilted this country's trade with Turkey in unusual ways towards gold (Financial Times, February 18, 2013). We also eliminate Syria, because that country's 2004 free trade agreement with Turkey that became effective on January 1, 2007 was suspended on December 6, 2011 (after the start of our sample) due to political differences over the handling of the Syrian civil unrest that turned into a civil war. As a result, 2012 bilateral trade between the two countries shrank by 74% down to $ 566 million compared to its 2011 level. After these exclusions, our sample corresponds to roughly 83.6% of the original dataset.

Long-term bank ratings by Moody's, S&P and Fitch are obtained from the Bankscope database. We collect all the available ratings from these three agencies for individual depository financial institutions, which 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.[28] Many banks we have multiple agency ratings, in which case we follow the rules imposed by the Turkish banking regulators (BDDK Directive, June 28, 2012, Supplement 1, Section 2, Articles 1.5 through 1.6). If a foreign counterparty has two agency ratings, Turkish banks have to use the worst (lower) of the two ratings. If a counterparty has three ratings, the banks are required to use "the better of the worst two ratings" (i.e., the middle rating). We obtain weighted-average ratings for each country for each annual period (pre- and post-July 1, 2012) using depository institutions' latest available total assets as well as the number of days the selected rating was valid (if there were changes to that institution's ratings over time). After dropping countries whose average bank rating proxy did not remain in the same risk-weight range (i.e., in a given row of Table 1) between July 1, 2011 and June 30, 2013, we are left with 3,828 observations for the OECD sample and 19,140 observations for the non-OECD sample. As a robustness check, we also use sovereign country long-term debt ratings (obtained from www.countryeconomy.com) as a separate proxy for bank ratings, subject to the same rules above.

For destination-country total imports we use the quarterly IMF Direction of Trade Statistics (DOTS) imports data by country of origin between July 1, 2011 and June 30, 2013. For each destination country *c* we aggregate quarterly imports from all other countries after excluding shipments from Turkey. Then, we match the quarterly country-industry Turkish exports data with the quarterly IMF country-level imports data before aggregating them into one pre- and one post-Basel II annual period. As a result of these restrictions we end up with exports to 160 countries along 22 ISIC industry sectors under three different methods of payments (CIA, CLC and OA) for two years centered on July 1, 2012. The number of countries in our time-varying country fixed effects

[28] 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.

regressions increases to 174 as we have 14 countries for which imports are not available in DOTS, yet these states have TUIK exports data that satisfy the restrictions that we impose.

Distance data between the capital cities for Turkey and export destination countries are obtained from the Centre d'Etudes Prospectives et d'Informations Internationales (CEPII) database. The indicator variable *D_ADJACENT* is coded as one for the four countries that have a land border with Turkey (which are the neighbors that remain in the dataset after the restrictions we need to impose: Armenia, Bulgaria, Georgia, and Iraq) and zero otherwise.[29]

The summary statistics are provided in Table 2 for OECD and non-OECD subsamples. The OECD sample contains 3,828 observations (= 29 countries × 22 ISIC industries × 3 payment terms × 2 annual periods), whereas the non-OECD sample contains 19,140 observations (= 145 countries × 22 ISIC industries × 3 payment terms × 2 annual periods). For the OECD (non-OECD) sample the dependent variable, *EXPORTS*, has a mean of $ 32.7 million ($ 6.0 million) and a standard deviation of $ 145 million ($ 42 million). Such large standard deviations are typical in international trade studies: *EXPORTS* to OECD (non-OECD) countries range from zero to $ 2.6 billion ($ 2.1 billion), with a median of $ 0.90 million ($ 20 thousand). In fact, for OECD (non-OECD) observations approximately 14.3% (37.5%) of country-industry-year observations in the dataset are equal to zero. The annual country-level imports excluding imports from Turkey (*IMPORTS_EX_TUR*) to the 29 OECD (131 non-OECD) countries has a mean of $ 367.7 billion ($ 49.6 billion). The average *DISTANCE* between the capital cities of OECD (non-OECD) export-destination countries and Turkey is 4.6 thousand (6.1 thousand) kilometers.

When looking at the distribution of ratings groups for the OECD sample, presuming that CLCs have remaining maturities longer than three months, we observe that 20.7% of the observations (including zero exports) belong to countries whose banks are rated, on average, between Aaa and Aa3 (the base-case range for OECD countries for which the RW remains 20%), while 79.3% to countries whose financial institutions have average ratings that range between A1 and Baa3 (for which the RW increases from 20% to 50%). For the non-OECD sample, assuming that CLCs have maturities longer than three months, we observe that only 2.1% of the observations belong to countries whose financial institutions are rated, on average, between Aaa to Aa3 (for which the RW decreases from 100% to 20%); 57.2% to countries whose banks have average ratings ranging from A1 to Baa3 or that are non-rated (for which the RW decreases from 100% to 50%); and the remaining 40.7% to countries whose CLC-issuing institutions are rated Ba1 to B3 (the base case, unchanged RW, category for non-OECD countries). A slightly different distribution is obtained for the non-OECD group when we assume that the risk-weights for CLCs with maturities lower than three months.

[29] Note that *D_ADJACENT* never appears in Tables 3 through 6. This is because (i) in the OECD case regressions reported in Tables 3 and 4, Greece, the only member-neighor, drops out as it moves from default to specualtive grade rating range during the period we examine, and (ii) in the non-OECD case we only report industry-time and country-time fixed-effects models in which *D_ADJACENT* becomes redundant.

Before discussing the estimation results, we go over some of the patterns in the data, which will help us in interpreting the model's estimates. We observe that the share of CLC-financed exports ranges around 12.1% throughout the pre-Basel II adoption period.[30] In the pre-Basel 2 period there is a difference between OECD and non-OECD countries in the usage of CLC-based instruments. CLC exports account for 6.4% of trade value towards OECD countries while 17.9% for non-OECD countries. In our data, bank financing (using CLCs) ranks the second after exporter financing (using OA) but comes before importer financing (using CIA), in terms of value.

There exists considerable heterogeneity in the use of CLCs across industries. In 2010, an out-of-sample year for our study, the share of CLC-based exports had an average of 8.1% and a standard deviation of 10.6% across the 22 two-digit ISIC industries we consider.[31] To illustrate, CLC-financed exports accounted for only 0.82% of exports in the manufacture of office, accounting and computing machinery, in contrast to 40.7% of exports in the manufacture basic metals. The example is consistent with the explanation provided by Antras and Foley (2013) for the small share of CLC-financed exports in the food industry: goods produced in the basic metal industry may be easier to collateralize than those produced in the manufacture of office, accounting and computing machinery. Another explanation might be related to the relative transaction sizes in the two industries. For large transactions, it is easier to cover the fees charged by banks when they issue CLCs for their importer clients or when clearing (accepting) them for their exporter clients. Since transaction sizes are expected to be much larger in basic metals industry than in the manufacture of office, accounting and computing machinery, it is not surprising to see wider use of CLCs in the former. Thus we are more likely to see an effect of the Basel II implementation on the use of CLCs in industries that have always relied more on this sort of financing. In section 4, we will test whether such heterogeneity across industries exists. Next we present the results of our empirical results.

4. Results

Before starting to discuss the results, we note that all of the OECD sample regressions are estimated with robust standard errors as the number of countries involved (24) may be too small, as Kezdi (2004) suggests that roughly 50 clusters are needed for clustered standard errors to be effective. In contrast, all of the non-OECD sample regressions in the paper are estimated with standard errors clustered at the country-level given the larger number of countries involved. In all of the PPML regressions we impose Stata's "strict" options, which prevents the problem of overfitting of the model

[30] This percentage is in the range of ratios of CLC-financed trade reported for a limited number of countries in the literature. Ahn (2013) states that 4% of Columbian imports between 2008-2009, and 20% of South Korean exports by 2012. Antras and Foley (2013), examining export transactions of a large producer of frozen poultry products in the US, find that 6% of its exports are CLC-based.

[31] We provide 2010 data on CLC prevalence because we use them to classify industries into high- and low-CLC usage as a part of our robustness checks.

to zero-trade observations (as a result, the number of observations may vary across columns for a given sample depending on the set of fixed effects included in the empirical model).

4.1. OECD-country sample estimates

First we examine the OECD sample assuming that the CLCs have remaining maturities higher than three-months on average, in which case the RW increase from 20% under Basel I to 50% (a 150% increase) for A1 to Baa3 rated (treated) counterparties after Basel II compared to Aaa to Aa3 rated (non-treated or base-case) export-destinations for which the RW remains at 20%. Our hypothesis is that we should see a decrease in exports for the treated group after Basel II adoption.

The basic results for the OECD sample are presented in Table 3. The OLS estimates of Eq. (3) are in columns A and B, and the PPML estimates of Eq. (4) are in columns C through F. The differences across the columns are due to different fixed effect combinations: there are industry fixed effects in columns A and C, separate industry and country fixed effects in columns B and D, separate industry and country-time fixed effects in column E, and separate industry-time and country-time fixed effects in column F.

In discussing Table 3 results, we first focus on the explanatory variables that are linked with the gravity models in the international trade. In this literature the coefficient estimates are found to be close to +1 for the logarithm of the destination country GDP and around -1 for the logarithm of distance. In column A (OLS estimates with industry fixed effects) we observe that the coefficient estimate for *ln(IMPORTS_EX_TUR)* is equal to 0.947 and in column C (PPML estimates with only industry fixed effects) it is equal to 1.030, both of which are statistically significant at the 1%-level: as the aggregate import demand of the destination country increases by 1%, Turkey's industry-level exports to that country increases by 0.95%, which is similar to the typical coefficient estimate of GDP in gravity models. In column, the coefficient estimate for *ln(DISTANCE)* is equal to -0.923 (statistically significant at the 1%-level). These coefficient estimates are close to the estimates observed in the trade literature. However, in column B where we add country fixed effects together with industry fixed effects to OLS models, the coefficient estimate for *ln(IMPORTS_EX_TUR)* is equal to 1.697 but not statistically significant. This is not too surprising as column B looks at *within* country variation over a short period of time (given the country fixed effects) whereas typically gravity regressions rely on cross-sectional or pooled-OLS estimators (over longer periods) that focus on *across* country variation (which is the case of the regression model of column A). Similar observations can be made for *ln(IMPORTS_EX_TUR)* and *ln(DISTANCE)* in columns C and D when we look at the estimates of Eq. (4) using PPML. We conclude that our pseudo gravity models yields

reasonable estimates for *ln(IMPORTS_EX_TUR)* and *ln(DISTANCE)* even though we do not estimate a full-blown gravity model.[32]

Next, we turn our attention to the difference-in-differences part of Eq. (3) in the log-linear specifications presented in columns A and B of Table 3. The coefficient estimates for *D_CLC* and *D_CIA* are negative whereas that for *D_BASELII* is positive (all of which are statistically significant at the conventional levels): OA-based exports are more prevalent than CLC- or CIA-based shipments, and OA-based shipments increase in the year after the adoption date on July 1, 2012, respectively. The finding that exporter (OA) financed exports are increasing after adoption is compatible with the following conjecture (for which we have no direct evidence): if Basel II leads to an increase in domestic lending to small and medium enterprises (SMEs) because the RW for loans to SMEs are reduced from 100% to 75% (NEED TO FIND ACTUAL REF AND MORE DETAIL IS NEEDED HERE), there could be a spillover effect from an increase in domestic lending due to Basel II generating a larger availability of working capital, which in turn leads to an increase exporter-financed OA shipments. In column A, the estimates for *D_CLC×D_BASELII* and *D_CIA×D_BASELII* are not statistically significant, something that holds when we also add country fixed effects in column B. The coefficient estimates for *D_A1-Baa3* and *D_CIA×D_A1-Baa3* are not statistically significant in columns A or B: prior to Basel II adoption there is no difference between flows to countries whose banks have average ratings between Aaa-Aa3 versus A1-Baa3 according to the OLS estimates when OA- and CIA-based financing are considered. The coefficient estimate for *D_CLC×D_A1-Baa3* is positive and statistically significant at the 1%-level: there is more CLC-financed trade to states whose banks have an average rating between A1-Baa3. The coefficient estimate for the interactions of *D_A1-Baa3×D_BASELII*, as well as those for *D_CIA×D_A1-Baa3*, *D_CIA×D_BASELII*, and *D_CIA×D_A1-Baa3×D_BASELII* are all statistically insignificant in columns A and B.

Importantly for us, the estimate for the triple interaction *D_CLC×D_A1-Baa3×D_BASELII*, which is the focus of our test, is negative but statistically insignificant in either column A or B of Table 3. At a first glance these results might suggest that Basel II induced risk-weight changes for to A1 to Baa3 rated counterparties had no impact on Turkey's CLC-based exports when the associated RW increased from 0.20 to 0.50. But, as Santos Silva and Tenreyro (2006, 2011) show, log-linear models (such as those presented in columns A and B), which do not take into account zero-trades and do not handle heteroskedasticity properly, might yield biased and inconsistent gravity model estimates. Since approximately roughly 14% of our OECD observations are nil, the omission of zero trade flows is likely to be an important source bias for the OLS estimator.

[32] This observation also holds true for the corresponding coefficients of equations (5) and (6) using the non-OECD sample (which are discussed under Section 4.2), for which we only report *D_CLC* and its interactions to conserve space. Full model estimates are available from the authors upon request.

In columns C and D of Table 3, we present the results of PPML estimates with all possible 3,828 OECD observations. Our focus is on the difference-in-differences part of the empirical models. In column C the coefficient estimate for *D_A1-Baa3* is equal to -0.303, which, in contrast to column A findings, is statistically significant at the 1%-level: in the pre-Basel II period OA-based exports to destinations whose banks had an average rating between A1 and Baa3 are 26% less likely (with an incidence ratio of -0.261 = $e^{-0.303}$ -1). Again, as in column A, the coefficient estimate for *D_BASELII*, which measures the change in OA exports to countries whose banks have ratings between Aaa and Aa3 on average, is positive but not statistically significant. These differences indicate that excluding zero-trade flows do indeed bias estimates. We find that the coefficient estimate for *D_A1-Baa3×D_BASELII* in column C changes sign compared to columns A and B but is still not statistically significant: in the year that follows Basel II there is no change in the export flows countries with average bank ratings in the A1 to Baa3 range. Similar to columns A and B, the coefficient estimate for *D_CLC* (*D_CIA*) is negative and statistically significant at the 1%-level: after accounting for industry and country fixed effects, in the year prior to Basel II adoption there is a 97.5% (= $e^{-3.657}$ – 1) lower incidence of CLC or CIA based trade compared to OA based trade to countries with banks rated Aaa through Aa3. This last result is robust to the addition of country-time (in column E) and country-time and industry-time fixed effects (in column F). We also note that in the pre-adoption year there is an economically and statistically significant increase in the incidence ratios for CLC- and CIA-financed exports for A1 through Baa3 counterparties (the corresponding coefficient estimates are 1.257 and 0.871, respectively).

Now we turn our attention to our test, i.e., the coefficient estimate of the triple interaction *D_CLC×D_A1-Baa3×D_BASELII*: in the PPML regression of column C the coefficient estimate is equal to -1.509, which is statistically significant at the 5%-level. This suggests that when we use the correct estimator that takes into account zero-trade observations (and the heteroskedasticity of trade flows by country of destination) the incidence of exports to countries whose banks have an average rating between A1 to Baa3 decreases by 78% (= $e^{-1.509}$ – 1) when the corresponding RW increases by 150% (from 20% to 50%). We can calculate the associated RW elasticity of CLC-exports as -0.52 (= -0.78/1.50), which suggests that a 1% increase in the capital charge for CLCs leads to a -0.5% drop in CLC-based exports to OECD countries with A1 to Baa3 rated banks. These results are robust to the addition of country (in column D), country-time (in column E), and country-time and industry-time fixed effects (in column F). The coefficient estimate for the triple interaction *D_CIA×D_A1-Baa3×D_BASELII* is again, as expected, statistically not different from zero in either of the columns.

In Table 4 we examine the robustness of the Table 3 results for the OECD sample, by re-estimating Eq. (4) with separate country-time and industry-time fixed effects using PPML (i.e., by using the empirical model in the last column of Table 3). To conserve space, we show only the coefficients of interest pertaining to CLC-based exports as the results of other coefficients do not

change qualitatively.[33] Our discussion of the results focuses on the triple interaction, i.e., on the test of our hypothesis. In column A (B) we show the estimates for industries that respectively had an above (below) median use of CLC-trade in total exports in the year 2010 (which is outside of the estimation period). The results indicate that the post-Basel II decrease in CLC-financed exports to countries with A1 through Baa3 rated banks on average (when the related RW increases from 20% to 50%) that we observe in Table 3 is driven by industries that rely on this payment type more: in column A, the estimate for *D_CLC×D_A1-Baa3×D_BASELII* is equal to -1.999 (statistically significant at the 5%-level), suggesting a drop in related exports by -86.5%. In contrast, the corresponding estimate in column B is equal to -0.0476 but not statistically significant. In column C we present the results of a winsorized PPML regression of Eq. (4) in which top 5% observations for each industry in each year are replaced by the value of the 95th percentile of their industry: the coefficient estimate for *D_CLC×D_A1-Baa3×D_BASELII* is equal to -1.148, which suggests a -68% change in the incidence of exports to destinations with A1 to Baa3 rated banks post-Basel II. The (approximately) 10% difference in the incidence rates implied by the coefficient estimates of *D_CLC×D_A1-Baa3×D_BASELII* in column F of Table 3 and column C of Table 4 suggests that some of observed effect is due to observations in the tail of the distribution. In column D of Table 4, Eq. (4) is reestimated using a square panel with strictly positive export values for all payment types and periods. *D_CLC×D_A1-Baa3×D_BASELII* coefficient estimate is equal to -1.515 (statistically significant at the 5%-level), which suggests that our findings are not driven by the extensive margin (i.e., by starting or stopping exports by Turkish exporters to different destinations): looking at a sample with only intensive margin (albeit only at the industry-country level and not at the firm level), we obtain very similar results compared to the last column of Table 3. In column E, we estimate Eq. (4) using quarterly data between July 1, 2011 and June 30, 2013 (with now industry-quarter and country-quarter fixed effects). The coefficient estimate for the triple interaction of interest is equal to -1.605 (statistically significant at the 5%-level). This finding suggests that the bias that might be due to autocorrelated errors (indicated for panel difference-in-differences models in Bertrand, Duflo, Mullanaithan, 2004) appears to be small. Finally, in the last column of Table 4, we conduct a placebo test using a fictitious Basel II adoption date of July 1, 2011 and exports data between July 1, 2010 and June 30, 2012 that are aggregated over two annual periods as before: the coefficient estimate for *D_CLC×D_A1-Baa3×D_BASELII* is equal to 0.0181, which is not statistically significant. This suggests that our findings are not due to some unexplained seasonality in the data. Next, we present the results of tests conducted with the non-OECD sample.

4.2. Non-OECD country sample estimates

[33] Full tables are available from the authors upon request.

In Table 5 we present pseudo gravity model estimates for the non-OECD sample assuming that the CLCs have on average maturities longer than three months. Given the bias exhibited by the log-linear models in Section 4.1, we only report PPML estimates of Eq.(5) with the most flexible fixed-effects combination (i.e., with separate time-varying industry and time-varying country fixed effects) to capture as much time varying observables to the extent country-industry-year data allows us to. Column A presents the results for the main sample, columns B through G present the robustness checks with different subsamples and report them in the same order as Table 4. To conserve space, we only present the estimates of *D_CLC* and its interactions. For the non-OECD sample our tests are based on RWs that decrease with Basel II adoption. As a result, now we expect the related exports to increase under Basel II.

First, we focus on column A of Table 5. Pre-Basel II the incidence of CLC-based exports to countries whose banks have speculative (but non-default) grade ratings on average are -73% (=$e^{-1.316}$-1) lower than OA-based exports to the same destinations. During the same period the incidence of CLC-financed shipments to non-OECD countries with banks rated Aaa to Aa3 is almost 500% (=$e^{1.791}$-1) higher, whereas those for destinations with A1 through Baa3 or non-rated financial institutions are not statistically significant. Post-adoption the incidence of CLC-based exports to countries whose banks have speculative (but non-default) grade ratings on average are -36% (=$e^{-0.453}$-1) lower, a result that is statistically significant at the 1%-level. Next, we turn our attention to our tests.

In column A, the triple interaction *D_CLC×D_Aaa-Aa3×D_BASELII* has a coefficient estimate of 0.601, whereas *D_CLC×D_A1-Baa3&NR×D_BASELII* has a coefficient estimate of 0.408, both of which are statistically significant at the conventional levels. These results are comforting in the sense that the larger (80%) drop in the RWs for higher investment grade (Aaa to Aa3) range (from 100% to 20%) is followed by higher increase in exports (0.601) compared to a smaller increase in exports (0.408) following a relatively smaller (50%) drop in RWs (from 100% to 50%) for the lower investment grade (A1 to Baa3) and not rated counterparties. The corresponding country-industry level export elasticities to RW changes are -1.03 (= [$e^{+0.601}$-1] / [-0.80]) for countries whose banks are rated Aaa-Aa3 on average and -1.01 (= [$e^{+0.408}$-1] / [-0.50]) for states whose banks are rated A1-Baa3 on average or are not rated. In other words, a 1% decrease in CLC-trade costs leads to an approximately 1% increase in CLC-related exports.

Next, we check the robustness of these results. In columns B and C of Table 5, we reestimate Eq. (5) using the subsample of industries that use CLCs higher or lower than sample average in 2010, respectively. We observe that in column B the coefficient estimates for *D_CLC×D_Aaa-Aa3×D_BASELII* and *D_CLC×D_A1-Baa3&NR×D_BASELII* are slightly higher (0.677 and 0.419, respectively, both of which are statistically significant at the 1%-level) when compared to their respective column A estimates (but we do not test whether the apparent difference between the two sets of coefficients is statistically significant). In contrast in column C the estimate for *D_CLC×D_Aaa-Aa3×D_BASELII* is equal to 1.247 (which is only marginally statistically significant), whereas

the estimate of *D_CLC×D_ A1-Baa3&NR×D_BASELII* is equal to 0.0440 but not statistically significant. We conclude that the estimates for the triple interactions in column A are driven by industries that rely more on CLC-financing. In column D, we check for the potential influence of outlier observations after winsorizing the non-OECD data at the 5th percentile of their distribution's right tail. Indeed, the estimates for *D_CLC×D_ Aaa-Aa3×D_BASELII* and *D_CLC×D_ A1-Baa3&NR×D_BASELII* are lower (0.201 and 0.271, respectively, but only the second one is statistically significant at the 5%-level). Using the estimate for *D_CLC×D_ A1-Baa3&NR×D_BASELII*, we find an export elasticity of -0.623 (= [$e^{+0.271}$-1] / [-0.50]). In column E, we examine whether our results are driven by the intensive versus extensive margin (albeit at the country-industry level). We find that a square-panel sample in which there are country-industry level exports in both years yields results that are very similar to those of column A: the coefficient estimates for *D_CLC×D_ Aaa-Aa3×D_BASELII* and *D_CLC×D_ A1-Baa3&NR×D_BASELII* are equal to 0.626 and 0.406, both of which are statistically significant at the 1%-level. In column F, we find similar results using quarterly data (and quarterly varying industry and country effects): the coefficient estimates for *D_CLC×D_Aaa-Aa3×D_BASELII* and *D_CLC×D_A1-Baa3&NR ×D_BASELII* are equal to 0.716 and 0.393, both of which are statistically significant at the conventional levels. Finally, the results of a placebo test, in which the Basel II adoption is fictitiously set equal to July 1, 2011, reveal no statistically significant effects for the triple interactions.

In Table 6, we estimate Eq. (6) assuming that the average export-CLC maturities are less than three months. In this case the identification rests on the fact that the RW decreased from 100% to 20% for counterparties with investment grade ratings (Aaa to Baa3) and those that are not rated, whereas the RW for counterparties with speculative (but non-default) ratings was reduced to 50%. As in Table 5, for the sake of brevity we only present the estimates of *D_CLC* and its interactions and focus on the coefficient estimates of the triple interaction *D_CLC×D_Aaa-Baa3&NR×D_BASELII*. Table 6 is structured in the same way as Table 5 (with the same set of time-varying fixed effects). In column A we present the results of Eq. (6) for the main sample. The coefficient estimate for *D_CLC×D_ Aaa-Baa3_&_NR×D_BASELII* is equal to 0.444, a result that is statistically significant at the 1%-level. As the Aaa to Baa3 rated or non-rated non-OECD counterparty RW decreases from 100% to 20%, the incidence of exports to countries whose banks are rated as such increases by 55.9% (= $e^{+0.444}$-1). In calculating the corresponding elasticity of exports to RW changes we need to take into account the fact that the RW associated with exports to Ba1 through B rated countries fall as well from 100% to 50%. Therefore, the elasticity is equal to -0.93% (= 0.559 / [(0.2-0.5) / 0.5]): a 1% decrease in CLC cost due to RW change leads to a 0.93% increase in exports.

Next, in columns B through G of Table 6, we conduct the same series of robustness checks as in the corresponding columns of Table 5. In column B we find that the results for the main sample are driven by industries with above median CLC-use (the estimate for *D_CLC×D_ Aaa-Baa3_&_NR×D_BASELII* is equal to 0.448, a result that is statistically significant at the 1%-level). In

column C, we find no such effect for industries with below median CLC-use (triple interaction's estimate is equal to 0.210 but not statistically significant). The estimate of elasticity is almost halved when we winsorize the largest 5% of observations by the corresponding 95th percentile value of the distribution for that industry: the export elasticity to RW changes becomes equal to -0.52 (= [$e^{0.272}$-1] / [(0.2-0.5)/0.5]). A 1% decrease in CLC-related costs (associated with RW changes) leads to a 0.5% increase in country-industry level exports. In colum E, we reestimate Eq. (6) with a square panel, and observe that the results are driven by the intensive margin (rather than the extensive margin) when the latter is defined at the country-industry (rather than firm) level: the coefficient estimate for the triple interaction is equal to 0.446, which is statistically significant at the 1%-level. In column F, we reestimate Eq.(6) with quarterly data and find very similar results to those of column A: the coefficient estimate for the triple interaction is equal to 0.429, which is statistically significant at the 1%-level. Finally, in column G we conduct a placebo test using the fictitious Basel II adoption date of July 1, 2011, and unsurprisingly we find no statistically significant effect.

Finally, we estimate equations (4) through (6) using long-term sovereign debt ratings as a proxy for export-destination countries' average bank ratings. In these estimates, presented in Appendix Table A2, we obtain similar coefficient estimates for the triple interaction compared to those presented in tables 3 through 6, but the coefficient estimates are lower and their statistical significances are weaker. For the OECD sample (column A of Table A2), the triple interaction coefficient estimate of -0.698 has the correct negative sign but is almost half the size of the comparable estimate reported in the last column of Table 3, and is not statistically significant. For the non-OECD sample (columns B and C of Table A2) triple interactions' coefficient estimates are somewhat smaller (0.385 and 0.305) than comparable estimates in the first columns of tables 5 and 6, but have weaker statistical significance levels. These results corraborate our earlier findings.

5. Interpretation and conclusion

In this study we make contributions to two different strands of research. First, we find that Basel II adoption has an impact on the real economy through exports (which are a component of the GDP). While the effect of adoption or changes in risk-based capital requirements on banks' supply of loans has been documented, we know of no *direct* evidence of their impact on the real sector besides Brun, Fraisse, and Thesmar (2013). These authors, use loan-level data and empirically show that Basel II's adoption affected French firms' borrowing, working capital, employment and investment in a positive way. We provide evidence that the impact of changes in risk-based capital requirements also affect aggregate trade flows.

Using industry-destination country trade data for Turkey, we document that Basel II adoption affects CLC-based shipments to different groups of countries differentially given the various changes in the risk-weights associated with banks' holding of CLCs. Importantly, we obtain consistent results with two different sets of countries (OECD and non-OECD) for which the identification schemes

differ (the related RW changes are in the opposite directions). While our set-up does not allow us to draw precisely the welfare impact of Basel II adoption we show that differential changes in capital charges for export-related CLCs, which are held by Turkish banks as off-balance sheet items, affect trade flows. Our findings are also consistent with Turkish banks rationing CLC-acceptance when it becomes relatively more costly for them to do so, though we cannot differentiate between these two stories. Our evidence is robust to various specifications and estimations with various subsamples of the data and we find no evidence with placebo tests.

Our second contribution is regarding the cost of trade-term elasticity of exports. These range between -0.5 to -1 approximately depending on the sample (OECD versus non-OECD) and the changes in the RW moving Basel I to Basel II (which range from 50% to 150%). The elasticities are lower for the OECD sample (approximately -0.5) for which the pre-Basel II usage of CLC is lower (6.4% of exports value while 17.9% for non-OECD countries). These elasticities range between -0.45 and -0.62 when we winsorize the values of exports at the industry level to guard against the influence of large export flows driving our results. These elasticities may in part be also due to the ongoing substitution between different payment terms used that we cannot observe. We stress that the impact we find pertains to a fraction of Turkey's exports and not to their totality.

If we assume that the full savings of holding CLCs due to RW decreases were passed by banks on to their exporter-customers and that the letters of credit cost typically 1%-5% of a trade transaction, these numbers would mean that the cost of exporting would drop by 0.5% to 5%. The more conservative estimates (based, for example, on the elasticities obtained using the samples winsorized at the 95th percentile of industry level export flows), would suggest a -0.46% to -0.62% change in exports given a 1% change in the cost of CLCs. We can also back-out what would be a trade fall in CLC-based Turkish imports of a country that would lose its good bank ratings. Relying for example on estimates of the treatment effect from the non-OECD sample shown in Table 6, if a country's banks would lose an investment grade rating on average (*Aaa-Baa3*) and pass to non-investment grade (*Ba1 to B3*) this would imply a 31.3% to 55.9% fall in the incidence of CLC-financed trade if all trade contracts were to have remaining-maturities that are shorter than three months (assuming that there are no other changes when moving between rating categories). Given the average usage of CLC instruments in the medium non-OECD investment category of 16.7% in total trade value (and assuming no substitution among the methods of payment) this implies a fall between 5.2% and 9.3% of total trade value (given triple interaction estimates of columns D and A of Table 6, respectively). Such a deterioration of ratings occurred for example to Bulgaria or Jordan in the period from 2010 to 2013. As a result, there may be not only a trade penalty for sovereign default as found in Rose (2005) but also a trade cost if the average ratings of banks in the counterparty country simply worsen or a sovereign downgrade impacts negatively the ratings of banks domiciled in that country.

To conclude, we find an economically important sensitivity of Turkish exports to changes in trade finance (as measured by our elasticity estimates) when trade is intermediated through the use of

CLCs. However, the effect on total trade is much weaker given the 6.4% (17.9%) share of CLC in Turkey's total exports to OECD (non-OECD) countries. In other words, we find that the trade-finance channel may play an important role in determining the export flows, similar to the findings of Amiti and Weinstein (2011) or Chor and Manova (2012). However, our emphasis is on the role of risk-based capital regulation for banks. How to reconcile that with the claims by Eaton et al. (2011) that the bulk of the fall of trade during the Great Recession could be attributed to a fall in trade in durable goods (which could have been affected in turn by a fall in demand)? If the patterns in Turkish trade generalize to other countries, perhaps this is because CLC instruments are often used in financing durable and investment good sectors (like machinery and equipment) or inputs into durable goods (like iron and steel) and changes in demand factors may coincide with the changes in financing terms.

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

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

Risk-Weights based on long-term agency ratings					
Basel I		Basel II			
Risk-Weights		Risk-Weights		Agency Rating Categories	
A	B	C	D	E	F
OECD	Non-OECD	CLC maturity < 3 months (in case original short-term ratings do not exist)	CLC 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)

Table 2
Summary statistics

This table presents the summary statistics for the data used in the analysis. *EXPORTS* (the dependent variable) is the value of exports (in millions of U.S. dollars) from Turkey to country *c* in industry *i* during quarter *t*. *IMPORTS_EX_TUR* is destination country *c*'s imports (in millions of U.S. dollars) from all other countries except Turkey during quarter *t* calculated using data from the IMF Direction of Trade Statistics (DOTS). *DISTANCE* is distance in kilometers between capital cities of destination countries and Turkey taken from CEPII database. Indicator variables, whose names are preceded by a prefix *D_*, are equal to one for if the observation belongs to the category of the indicator variable, and zero otherwise. *ADJACENT* stands for the eight countries that have a land border with Turkey; *CIA* stands for Commercial Cash in Advance; *CLC* for Commercial Letter of Credit; *OA* for Open Account. Please refer to Table 2 for Moody's (and the corresponding S&P and Fitch's) rating classifications.

PANEL A: OECD SAMPLE

Number of observations =3,828 (=29 countries ×22 industries ×3 payment terms ×2 annual periods)

	Risk-Weight	N	Mean	Median	Std. Dev.	Min.	Max.
EXPORTS		3,828	32.7	0.9	145.2	0	2,562.0
IMPORTS_EX_TUR		3,828	367,716.1	190,486.2	460,659.1	4,672.8	2,325,581.0
DISTANCE		3,828	4,579.8	2,209.7	4,507.6	1,123.0	1,7234.5
D_ADJACENT		3,828	0.000	0	0	0	0
D_CIA		3,828	0.333	0	0.472	0	1
D_CLC		3,828	0.333	0	0.472	0	1
D_OA		3,828	0.333	0	0.472	0	1
For CLC maturity > 3 months							
D_Aaa-Aa3	0.20	3,828	0.207	0	0.405	0	1
D_A1-Baa3	0.50	3,828	0.793	1	0.405	0	1

PANEL B: Non-OECD SAMPLE

Number of observations =19,140 (=145 countries ×22 industries ×3 payment terms ×2 annual periods)

	Risk-Weight	N	Mean	Median	Std. Dev.	Min.	Max.
EXPORTS		19,140	6.0	0.02	42.3	0	2,132.6
IMPORTS_EX_TUR		17,292	49,561.5	8,624.6	176,453.1	100.739	1,873,147.0
DISTANCE		19,140	6,119.0	5,433.4	3,855.0	442.1	16,859.5
D_ADJACENT		19,140	0.034	0	0.183	0	1
D_CIA		19,140	0.333	0	0.471	0	1
D_CLC		19,140	0.333	0	0.471	0	1

D_OA		19,140	0.333	0	0.471	0	1
For CLC maturity < 3 months							
D_Aaa-Baa3_&_NR	0.20	19,140	0.593	1	0.491	0	1
For CLC maturity > 3 months							
D_Aaa-Aa3	0.20	19,140	0.021	0	0.142	0	1
D_A1-Baa3&NR	0.50	19,140	0.572	1	0.495	0	1

Table 3
Log-linear and PPML regressions with annual country-and-two-digit-industry-level data: OECD sample, total assets weighted bank ratings.

This table presents log-linear fixed-effects and PPML fixed effects regression results with annual (pre- and post July 1, 2012) Turkish exports to OECD countries at the country-and-two-digit-industry-level using total assets weighted bank ratings. In columns (A) and (B) the dependent variable is the natural logarithm of $EXPORTS_{c,i,t}$ from Turkey to country *c* in industry *i* in period *t* in U.S. dollars [$ln(EXPORTS_{c,t})$]; in columns C, D, E and F it is $EXPORTS_{c,i,t}$. $ln(IMPORTS_EX_TUR_{c,t})$ is the natural logarithm of imports of destination country *c* in period t after excluding country *c*'s Turkish imports. $ln(DISTANCE_c)$ is the natural logarithm of the distance between the capital cities of the destination countries and Turkey. Indicator variables, whose names are preceded by a prefix *D_*, are equal to one for if the observation belongs to the category of the indicator variable, and zero otherwise. *BASELII* denotes after July 1, 2012 period following Basel II adoption in Turkey; *CLC* denotes commercial letters of credit-based exports; *CIA* denotes cash in advance-based exports; *A1-Baa3* for Moody's (S&P and Fitch) lower investment-grade risk-rating categories between A1 and Baa3 (A+ and BBB-) destination countries for which the Basel II risk-weight for CLCs is 0.50. For any payment type the omitted category is destination countries with Moody's ratings between Aaa and Aa3 (AAA and AA-) for which the Basel II risk-weight remains equal to Basel I risk-weight of 20%. The base case is formed by open account (OA) based exports. The regressions are with robust standard errors. Standard errors are provided within parentheses below the coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

	Log-linear (OLS) models				Poisson models							
	A		B		C		D		E		F	
ln(IMPORTS_EX_TUR)	0.947	***	1.697		1.030	***	3.122					
	(0.0268)		(1.632)		(0.0525)		(1.955)					
ln(DISTANCE)	-0.923	***			-0.967	***						
	(0.0433)				(0.0837)							
D_A1-Baa3	-0.130				-0.303	***						
	(0.194)				(0.110)							
D_BASELII	0.444	*	0.467	**	0.0352		0.122					
	(0.242)		(0.232)		(0.118)		(0.141)					
D_A1-Baa3×D_BASELII	-0.318		-0.321		0.0449		0.0508					
	(0.271)		(0.254)		(0.151)		(0.137)					
D_CLC	-4.302	***	-4.339	***	-3.657	***	-3.657	***	-3.608	***	-3.610	***
	(0.273)		(0.256)		(0.360)		(0.364)		(0.366)		(0.364)	
D_CLC×D_A1-Baa3	0.816	***	0.722	**	1.257	***	1.257	***	1.270	***	1.271	***
	(0.305)		(0.286)		(0.428)		(0.429)		(0.432)		(0.431)	

D_CLC×D_BASELII		0.0518		0.0856		1.583	**	1.583	**	1.637	**	1.638	**
		(0.389)		(0.364)		(0.705)		(0.702)		(0.702)		(0.701)	
D_CLC×D_A1-Baa3×D_BASELII		-0.0721		-0.0836		-1.509	**	-1.509	**	-1.578	**	-1.577	**
		(0.434)		(0.406)		(0.767)		(0.762)		(0.763)		(0.762)	
D_CIA		-2.717	***	-2.713	***	-3.646	***	-3.646	***	-3.646	***	-3.646	***
		(0.243)		(0.228)		(0.215)		(0.216)		(0.221)		(0.222)	
D_CIA×D_A1-Baa3		-0.0685		-0.0900		0.871	***	0.871	***	0.872	***	0.872	***
		(0.273)		(0.255)		(0.335)		(0.333)		(0.336)		(0.337)	
D_CIA×D_BASELII		-0.100		-0.102		0.551	*	0.551	*	0.555	*	0.555	*
		(0.345)		(0.323)		(0.333)		(0.332)		(0.331)		(0.331)	
D_CIA×D_A1-Baa3×D_BASELII		0.0160		0.0249		-0.681		-0.681		-0.684		-0.684	
		(0.387)		(0.362)		(0.462)		(0.456)		(0.456)		(0.456)	
Estimator		OLS		OLS		PPML		PPML		PPML		PPML	
Number of observations		3,278		3,278		3,828		3,828		3,278		3,278	
R^2		0.646		0.692									
Robust standard errors		yes		yes		yes		yes		yes		yes	
Fixed-Effects	*Industry*	yes		yes		yes		yes		yes		no	
	Country	no		yes		no		yes		no		no	
	Industry×time	no		no		no		no		no		yes	
	Country×time	no		no		no		no		yes		yes	

Table 4
PPML regressions with annual country-and-two-digit-industry-level data: OECD sample, total assets weighted bank ratings, robustness checks

This table presents PPML fixed effects regression results with Turkish exports to OECD countries at the country-and-two-digit-industry-level using total assets weighted bank ratings for the coefficients of interest (CLC-based trade) with other coefficient estimates suppressed to conserve space. Exports are aggregated at the annual (pre- and post July 1, 2012) level except for column E where they are at the quarterly level. Column A shows the results for the subsample with industries that used CLCs in exports above the industry median in 2010; column B gives the results for industries with industries that used CLCs in exports below the industry median in 2010; column C presents the results of a one-sided winsorized regression in which observations above the 95th percentile for each industry are replaced by the value at the 95th percentile of their industry; column D exhibits the results of a regression on "square" panel with all exports for a given industry-payment positive; column E displays the results with the exports aggregated quarterly; column F shows the results of a placebo regression where it is assumed that the BASEL II reform would occur fictitiously on July 2011 and exports are aggregated at the annual (pre- and post July 1, 2011) level. The dependent variable is the $EXPORTS_{c,i,t}$ from Turkey to country *c* in industry *i* in period *t* in U.S. dollars. Indicator variables, whose names are preceded by a prefix *D_*, are equal to one for if the observation belongs to the category of the indicator variable, and zero otherwise. *BASELII* denotes after July 1, 2012 period following Basel II adoption in Turkey (July 1, 2011 for column F); *CLC* denotes commercial letters of credit-based exports; *A1_Baa3* denotes lower investment-grade risk-rating categories between A1 and Baa3 by Moody's (A+ to BBB- by S&P and Fitch) for which the Basel II RW for export-CLCs is 0.50. For any payment type the omitted category is destination countries with Moody's ratings between Aaa and Aa3 (AAA and AA-) for which the Basel II risk-weight remains equal to Basel I risk-weight of 20%. The base case is formed by open account (OA) based exports The regressions are with robust standard errors. Standard errors are provided within parentheses below the coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

		A		B		C		D		E		F	
D_CLC		-3.471	***	-3.789	***	-2.783	***	-3.593	***	-3.628	***	-3.426	***
		(0.591)		(0.190)		(0.346)		(0.369)		(0.237)		(0.297)	
D_CLC×D_A1-Baa3		1.588	**	0.515		0.633		1.247	***	1.258	***	1.253	***
		(0.646)		(0.349)		(0.400)		(0.437)		(0.269)		(0.375)	
D_CLC×D_BASELII		2.078	**	-0.0217		1.219	**	1.586	**	1.730	***	-0.185	
		(0.860)		(0.521)		(0.585)		(0.702)		(0.615)		(0.470)	
D_CLC×D_A1-Baa3×D_BASELII		-1.999	**	-0.0476		-1.148	*	-1.515	**	-1.605	**	0.0181	
		(0.922)		(0.657)		(0.642)		(0.765)		(0.641)		(0.571)	
Estimator		PPML		PPML		PPML		PPML		PPML		PPML	
Number of observations		1,741		1,537		3,278		2,232		13,801		3,259	
Robust standard errors		yes		yes		yes		yes		yes		yes	
Fixed-Effects	*Industry×time*	yes		yes		yes		yes		yes		yes	
	Country×time	yes		yes		yes		yes		yes		yes	

Table 5

PPML regressions with annual country-and-two-digit-industry-level data: non-OECD sample, total assets weighted bank ratings, long-term CLC identification scheme; main results and robustness checks

This table presents PPML fixed effects regression results with Turkish exports to non-OECD countries at the country-and-two-digit-industry-level for the coefficients of interest (CLC-based trade) with other coefficient estimates suppressed to conserve space. The rating for each country was obtained by using total assets weighted bank ratings; the assumption that CLCs had maturities above three months was used. Exports are aggregated at the annual (pre- and post July 1, 2012) level except for column F where they are at the quarterly level. Column A presents the results for the main sample; column B for the subsample with industries that used CLCs in exports above the industry median in 2010; column C for industries with industries that used CLCs in exports below the industry median in 2010; column D for the one-sided winsorized sample in which observations above the 95th percentile for each industry are replaced by the value at the 95th percentile of their industry; column E for the "square" panel with all exports for a given industry-payment positive; column F for the quarterly aggregated exports sample; column G shows the results of a placebo regression where it is assumed that the BASEL II reform would occur fictitiously on July 2011 and exports are aggregated at the annual (pre- and post July 1, 2011) level. The dependent variable is the $EXPORTS_{c,i,t}$ from Turkey to country *c* in industry *i* in period *t* in U.S. dollars. Indicator variables, whose names are preceded by a prefix *D_*, are equal to one for if the observation belongs to the category of the indicator variable, and zero otherwise. *BASELII* denotes after July 1, 2012 period following Basel II adoption in Turkey (July 1, 2011 for column F); *CLC* denotes commercial letters of credit-based exports; *Aaa-Aa3* stands for higher investment grade risk rating categories between Aaa and Aa3 by Moody's (AAA to AA- by S&P and Fitch) for which the Basel II RW for export-CLCs is 0.20; *A1-Baa3_&_NR* denotes lower investment-grade risk-rating categories between A1 and Baa3 by Moody's (A+ to BBB- by S&P and Fitch) and countries for which there were no banks with a rating for which the Basel II RW for export-CLCs is 0.50. For any payment type the omitted category is destination countries with Moody's ratings between Ba1 and B3 (BB+ to B-) for which the Basel II risk-weight remains equal to Basel I risk-weight of 100%. The base case is formed by open account (OA) based exports. Regression standard errors are clustered at the destination country level. Standard errors are provided within parentheses below the coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

	A		B		C		D		E		F		G	
D_CLC	-1.316	***	-1.046	***	-3.349	***	-1.112	***	-1.377	***	-1.444	***	-1.559	***
	(0.268)		(0.254)		(0.227)		(0.191)		(0.276)		(0.276)		(0.241)	
D_CLC×D_Aaa-Aa3	1.791	***	1.894	***	1.248	***	0.945	***	1.897	***	1.763	***	2.036	***
	(0.365)		(0.298)		(0.385)		(0.289)		(0.393)		(0.375)		(0.388)	
D_CLC×D_A1-Baa3_&_NR	-0.0900		-0.0908		0.379		0.296		-0.0426		-0.0441		0.170	
	(0.477)		(0.478)		(0.527)		(0.244)		(0.489)		(0.478)		(0.461)	
D_CLC×D_BASELII	-0.453	***	-0.471	***	0.154		-0.187	*	-0.462	***	-0.392	***	0.115	
	(0.0874)		(0.0921)		(0.127)		(0.0965)		(0.0777)		(0.0881)		(0.138)	
D_CLC×D_Aaa-Aa3×D_BASELII	0.601	**	0.677	***	1.247	*	0.201		0.626	***	0.716	**	-0.235	
	(0.299)		(0.231)		(0.704)		(0.359)		(0.231)		(0.304)		(0.254)	
D_CLC×D_A1-Baa3_&_NR×D_BASELII	0.408	***	0.419	***	0.0440		0.271	**	0.406	***	0.393	***	-0.214	
	(0.131)		(0.139)		(0.166)		(0.108)		(0.128)		(0.136)		(0.183)	

Estimator		PPML	PPML	PPML	PPML	PPML	PPML	PPML
Number of observations		11,958	6,626	5,332	11,958	6,558	70,271	15,182
Clustered standard errors (country-level)		yes	yes	yes	yes	yes	yes	yes
Fixed-Effects	*Industry×time*	yes	yes	yes	yes	yes	no	yes
	Country×time	yes	yes	yes	yes	yes	yes	yes

Table 6

PPML regressions with annual country-and-two-digit-industry-level data: non-OECD sample, total assets weighted bank ratings, short-term CLC identification scheme; main results and robustness checks

This table presents PPML fixed effects regression results with Turkish exports to non-OECD countries at the country-and-two-digit-industry-level for the coefficients of interest (CLC-based trade) with other coefficient estimates suppressed to conserve space. The rating for each country was obtained by using total assets weighted bank ratings. The assumption is that CLCs have maturities below three months. Exports are aggregated at the annual (pre- and post Basel II adoption) level except for column F where they are at the quarterly level. Column A presents the results for the main sample; column B for the subsample with industries that used CLCs in exports above the industry median in 2010; column C for industries with industries that used CLCs in exports below the industry median in 2010; column D for the one-sided winsorized sample in which observations above the 95th percentile for each industry are replaced by the value at the 95th percentile of their industry; column E for the "square" panel with all exports for a given industry-payment positive; column F for the quarterly aggregated exports sample; column G shows the results of a placebo regression where it is assumed that the BASEL II reform would occur fictitiously on July 2011 and exports are aggregated at the annual (pre- and post July 1, 2011) level. The dependent variable is the $EXPORTS_{c,i,t}$ from Turkey to country *c* in industry *i* in period *t* in U.S. dollars. Indicator variables, whose names are preceded by a prefix *D_*, are equal to one for if the observation belongs to the category of the indicator variable, and zero otherwise. *BASELII* denotes after July 1, 2012 period following Basel II adoption in Turkey (July 1, 2011 for column F); *CLC* denotes commercial letters of credit-based exports; *Aaa-Baa3_&_NR* stands for investment grade risk rating categories between Aaa and Baa3 by Moody's (AAA to BBB- by S&P and Fitch) and countries for which there were no banks with a rating for which the Basel II RW for export-CLCs is 0.20. For any payment type the omitted category is destination countries with Moody's ratings between Ba1 and B3 (BB+ to B-) for which the Basel II risk-weight is 0.50. The base case is formed by open account (OA) based exports. Regression standard errors are clustered at the destination country level. Standard errors are provided within parentheses below the coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

	A		B		C		D		E		F		G	
D_CLC	-1.316	***	-1.046	***	-3.349	***	-1.112	***	-1.377	***	-1.364	***	-1.559	***
	(0.267)		(0.254)		(0.227)		(0.191)		(0.276)		(0.266)		(0.241)	
D_CLC×D_Aaa-Baa3_&_NR	0.0429		0.0482		0.413		0.334		0.0893		0.0508		0.316	
	(0.464)		(0.465)		(0.511)		(0.239)		(0.476)		(0.460)		(0.448)	
D_CLC×D_BASELII	-0.453	***	-0.471	***	0.153		-0.187	*	-0.462	***	-0.319	***	0.115	
	(0.0874)		(0.0921)		(0.127)		(0.0965)		(0.0777)		(0.0917)		(0.138)	
D_CLC×D_Aaa-Baa3_&_NR×D_BASELII	0.444	***	0.448	***	0.210		0.272	**	0.446	***	0.429	***	-0.232	
	(0.142)		(0.146)		(0.260)		(0.111)		(0.138)		(0.153)		(0.173)	
Estimator	PPML		PPML		PPML		PPML		PPML		PPML		PPML	
Number of observations	11,958		6,626		5,332		11,958		6,558		56,824		15,128	
Clustered standard errors (country-level)	yes		yes		yes		yes		yes		yes		yes	
Fixed-Effects *Industry×time*	yes		yes		yes		yes		yes		yes		yes	
Country×time	yes		yes		yes		yes		yes		yes		yes	

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

This table presents the Basel I and Basel II credit conversion factors (CCFs) for off-balance sheet commercial letters of credit (CLCs) 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).

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 (CCF)**	**Commercial Letter of Credit (CLC) Type**
5 (1) a) 3)	5 (3) a) 3)	100%	Export-related confirmed-CLCs
5 (1) b) 6)	5 (3) b) 5)	50%	CLCs that do not have a CCF of 100%, 20% or 0%
5 (1) c) 2)	5 (3) c) 2)	20%	CLCs 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 CLCs that do not require a payment to the recepient [exporter]

Table A2
Log-linear and PPML regressions with annual country-and-two-digit-industry-level data: sovereign ratings

This table presents PPML fixed-effects regression results with annual (pre- and post July 1, 2012) Turkish exports at the country-and-two-digit-industry-level using long-term sovereign credit ratings to classify export destinations to different risk-weight (RW) categories. The dependent variable is $EXPORTS_{c,i,t}$ from Turkey to country *c* in industry *i* in period *t* in U.S. dollars. Indicator variables, whose names are preceded by a prefix *D_*, are equal to one if the observation belongs to the category of the indicator variable, and zero otherwise. *BASELII* denotes period following Basel II adoption in Turkey after July 1, 2012; and *CLC* denotes commercial letters of credit-based exports. In column A, *A1-Baa3* denotes rating categories for which the Basel II RW for CLCs increases from 0.20 to 0.50, the base category RW remains at 0.20; in column B, *Aaa-Aa3* (*A1-Baa3&NR*) denotes credit rating categories for which the RW for CLCs decreases from 1.00 to 0.20 (from 1.00 to 0.50), the base category RW remains at 1.00; and in column C, *Aaa-Baa3&NR* denotes rating categories for which the RW for CLCs decreases from 1.00 to 0.20, the base category RW moves from 1.00 to 0.50 after adoption. The table only reports coefficient estimates of interest, those that correspond to open account and cash in advance categories are omitted for the sake of brevity. Standard errors, which are robust in column A and clustered in columns B and C, are provided within parentheses below the coefficient estimates. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

	A		B		C	
	OECD sample		Non-OECD sample		Non-OECD sample	
	CLC maturity > 3 months		CLC maturity > 3 months		CLC maturity < 3 months	
D_CLC	-2.915	***	-1.155	***	-1.155	***
	(0.295)		(0.339)		(0.339)	
D_CLC×D_BASELII	0.553		-0.353	**	-0.353	**
	(0.437)		(0.149)		(0.149)	
D_CLC×D_A1-Baa3	1.083					
	(0.729)					
D_CLC×D_A1-Baa3×D_BASELII	-0.698					
	(0.992)					
D_CLC×D_Aaa-Aa3			1.113	***		
			(0.359)			
D_CLC×D_A1-Baa3&NR			-0.486			
			(0.499)			
D_CLC×D_Aaa-Aa3×D_BASELII			0.155			
			(0.208)			
D_CLC×D_A1-Baa3&NR×D_BASELII			0.385	**		
			(0.176)			
D_CLC×D_Aaa-Baa3_&_NR					-0.108	
					(0.502)	
D_CLC×D_Aaa-Baa3_&_NR×D_BASELII					0.305	*
					(0.183)	
Estimator	PPML		PPML		PPML	
Number of observations	2,643		10,964		10,964	
Robust standard errors	yes		no		no	
Clustered std. errors (country-level)	no		yes		yes	
Fixed-Effects *Industry×time*	yes		yes		yes	
Country×time	yes		yes		yes	