Risk-Based Capital Requirements for Banks and International Trade

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Abstract

We provide the first evidence that changes in risk-based capital requirements for banks affect the real economy through international trade. Using a natural experiment – mandatory Basel II adoption in its Standardized Approach by all banks in Turkey on July 1, 2012 – we investigate the impact of new risk-weights applied to commercial letters of credit (CLC) on that country's exports to 174 countries. We estimate the resulting payment-term-cost elasticity of CLC-financed trade to be between -0.5 and -1 while the overall trade elasticity to be between -0.032 and -0.179. Calculations suggest that both CLC-related bank pricing and rationing channels are involved. (100 words)

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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 risk-based capital (RBC) requirements proposed by the Basel Committee on Bank Regulation of the Bank for International Settlements (BIS). For example, in 2009 Robert Zoellick, the then president of the World Bank, suggested that 10%-15% of the decrease in global trade during the Great Recession might be due to lower provision of trade finance under Basel II (Financial Times, February 19, 2009).[1] A 2009 survey by the International Chamber of Commerce (ICC) reports that "the feedback … on Basel II … [suggests] that most banks are facing tougher capital requirements for their [international] trade assets" (ICC, March 31, 2009, p. 40). Other banking surveys indicate that (i) Basel II had a negative impact on banks' provision of trade finance for 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, during its Seoul Summit the G20 stated that it would "… evaluate impact of regulatory regimes on trade finance" (G20, 2010). Four years later we still know very little regarding the impact of changes in capital standards on international trade due to scant research in this area.

To the best of our knowledge, we are the first to investigate the impact of regulatory changes in 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 requirements 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

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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 CLCs (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).

payment-default. The exporter presents the CLC for payment to its local (in our case Turkish) bank. If the bank, which charges a price for this service, accepts to clear the CLC for the exporter, it has to hold the CLC as an off-balance sheet item during the remaining maturity of the contract.[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 conversion factor* (CCF) that multiplies the nominal value of the CLC. Then, the capital requirement for 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 the 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 counterparty banks issuing export-related CLCs were domiciled in an OECD country or not. For the former the RW was 20% and for the latter it was 100%. 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. Here we use one of these three cases to illustrate how we identify the impact of Basel II on exports (more detail is provided in Section 3.1): for counterparty banks located in OECD countries and for CLCs with maturities longer 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 counterparty banks, or (ii) stayed constant at 20% for higher investment grade (Aaa to Aa3, or equivalently AAA to AA-)

[2] In some countries, such as the US, the confirmed CLCs can be sold in the money 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.

[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 Appendix Table A1 and the discussion in Section 3.1 for more details).

rated counterparty-banks, which form the base-case in our difference-in-differences regressions.[6] In this particular case, we hypothesize that Turkish exports involving A1 to Baa3 rated counterparty banks to decrease as the related RW (hence the capital requirement) increases. Our empirical results are in support of this hypothesis for the OECD sample. We find comparable results with the two other identification schemes that apply to the non-OECD sample.

We believe that our empirical findings are important given that there is little *direct* evidence on the effects of mandatory 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). There are two exceptions. Brun, Fraisse, and Thesmar (2013) 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% on average. Lee and Stebunovs (2012) examine the impact of changes in state-level bank capital ratios on average firm-size and firm creation in different industries following different capital regulations in the US during 1980s and 1990s (including, among others, the two-tiered adoption of Basel I RBC requirements partially in 1989 and fully in 1992). They find a negative impact of higher capital ratios at the state-level on average establishment size (as measured by the number of employees) but do not detect an effect on net firm creation. One of our main contributions is to expand this emerging line of research in a new dimension by evaluating the effects of Basel II on exports, which are, for many countries, an important part of the gross domestic product (GDP).

We also contribute to the emerging literature on the export or import payment terms and international trade flows (e.g., Antras and Foley, 2014; Auboin and Engemann, 2012; Glady and Potin, 2011; Mateut, 2012; Niepmann and Schmidt-Eisenlohr, 2013a and 2013b; Schmidt-Eisenlohr, 2013).

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6 In fact, three other categories were also possible post-Basel II for CLCs with maturities higher than three months issued by OECD banks: 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 need to impose on the data for a proper difference-in-differences estimation.

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 test the impact of Basel II adoption on exports that are based on these three different trade payment terms. Reassuringly, we find that Basel II adoption has no impact on CIA- and OA-based exports. 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; Chor and Manova, 2012; Paravisini, et al., 2014) 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 CLC pricing and rationing channels (unfortunately, our data do not allow us to distinguish between these two). Second, our findings suggest that changes in capital requirements that took place during the Great Recession might also have impacted CLC-based exports. [7] This is because as banks, hence their agency ratings, weakened, the trade punishment that Basel II incorporates through higher RWs increased.

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. 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).[8] FIRB

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

[8] Turkish banking authorities made it clear that the 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.

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 of SA by the Turkish banking regulators to all banks under their jurisdiction provides us with two sets of identification schemes that apply to two different (mutually exclusive) data samples and that imply opposite signs for the export flows.

Second, we work with data collected by the Turkish Ministry of Customs and Trade and provided to us by the Turkish Statistics Institute (TSI). 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.[9] 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.[10] 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

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[9] 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).

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

have to impose on the data (in order to be able to properly estimate difference-in-differences models), we still account for roughly 87% 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 (see Section 3.2.1 for more details). To control for time-varying country demand for foreign manufactured goods, some of our regressions incorporate country-level total imports for the period after excluding Turkish exports to that country. We also include distance between Turkey and the destination country as well as an indicator variable for adjacent countries to account for time invariant impediments to trade. We also estimate alternative specifications with country-time and industry-time fixed effects to account for unobservable variables. Inherent heteroskedasticity of trade flows to different destinations and/or omitted zero-trade flows lead to biased log-linear estimates. We rely on the Poisson Pseudo-Maximum Likelihood (PPML) estimator to deal with these problems as suggested by Santos-Silva and Tenreyro, 2006, 2011.

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 provided by Turkish banks (rather than shifts in the demand for such services by Turkish exporters). We should also note that our tests are inherently weak and potentially biased against us finding an effect as they rely on the joint hypotheses: (i) that our proxies for counterparty bank ratings are representative of the average of actual bank ratings involved in CLC-based exports to a certain country, and (ii) that Turkish banks react to Basel II-related differential changes in capital requirements (either through pricing or rationing 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 this 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 related RW decrease. In contrast, we do not observe similar patterns for exports settled under other methods of payment (i.e., CIA or OA). We can calculate the elasticity of CLC-intermediated exports to the changes in the RWs to be between -0.5 and -1.0. These findings suggest that the reaction of trade flows (in value) 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 substitution effects (across payment terms) in shipments, we estimate that the overall elasticity of total 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 roughly 0.03% to 0.18% decrease in trade flows. These estimates are comparable to those found Paravisini et al. (2014) 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 non-OECD countries with speculative-grade (Ba1 to B3, or equivalently BB+ to B-) ratings decrease post Basel II adoption. This finding is consistent with Turkish banks' imposing internal credit exposure limits for non-investment grade counterparties after Basel II as part of their risk-control management. These results are robust to the use of different proxies for counterparty ratings involved in CLC-financed exports to different destinations, the presence of zero-trade observations, frequency of the data (annual or quarterly), various sub-samples (sectors that rely more or less on CLC-financing in 2010), 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 empirical results (including robustness checks), whose economic significance is discussed 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. In contrast to the above papers that focus on the supply of loans by banks, two recent papers test for the effects of capital requirements on the real sector. Lee and Stebunovs (2012) use state-level US data to analyze the effect of increases in actual bank capital ratios, following various regulatory changes in the US, on firm size and creation in the manufacturing sector. After controlling for state-level branching deregulations, banking sector concentration, and demand-side factors, Lee and Stebunovs (2012) find that increases in state-level bank capital ratios lead to contractions in firm size (as measured by the number of employees) but have no effect on net firm creation at the state-level. Brun, Fraisse and Thesmar (2013) use much more detailed bank-and-firm matched loan-level French data to assess the impact of Basel II adoption in 2008 on bank lending as well as on corporate activity. 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. Controlling these as well as for unobservable bank heterogeneity and unobservable changes in firm credit demand, these authors 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, post-Basel II borrowing by French firms went up by roughly 12 billion euros (a 1.5% increase), whereas investment increased by 0.5%, leading to the preservation 235,000 jobs, 1% of the aggregate French employment. Our contribution is to expand this latest strand of literature by providing novel evidence that Basel II adoption also affects the real sector through an altogether different channel, i.e., via 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 from Peru, Paravisini et al. (2014) 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. 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. 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.[11] We complement this literature by examining the effects of a regulatory change that affects the capital charge (the implicit cost) of holding CLCs by

[11] That said, Levchenko, Lewis and Tesar (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.

Turkish banks. Changes in RWs due to mandatory Basel II adoption should affect these banks' pricing of export-CLCs, which in turn might affect exporters' behavior when serving different destination countries. Turkish banks might also ration holding higher risk export-CLCs for risk-management reasons, or might stop such rationning. The sizes of our coefficient estimates and a simple calibration exercise suggest that both channels are at play.

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. In contrast to the larger literature on (domestic) trade-finance (e.g., Giannetti, Burkart, and Ellingsen, 2011; and Klapper, Laeven, and Rajan, 2012), less is known on international trade-finance choices. A number of papers in this line of research focus on one type of international-trade payment at a time. For example, Mateut (2012) finds that CIA payments, which can be linked with both firm and industry characteristics, are used to reduce default risk in international trade. Glady and Potin (2011) provide a model in which asymmetric information and difficulties in contract enforcement increase CLC default risk that financial intermediaries are able to reduce. The key features of their model are supported in the data. 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- and CLC-based trade) 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 (2014) 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). Whereas aggregate data at our disposal do not allow us to examine trade finance trade-offs of exporters, as Antras and Foley (2014) do, 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 we can (i) account for changes in trade flows in a comprehensive way, and (ii) provide tests as to whether changes in the (implicit) cost of export-related CLCs affect shipments that are based on other payment terms (they should not given our hypotheses). In the next section, we detail the identification schemes, the empirical specifications and the data that we use.

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 capital that is needed to be set aside, which is equal to the CLC's nominal value times CCF times RW. When examining these changes our focus is on the changes that Basel II implies for RWs and treat CCFs as constant parameters. One reason for this focus on RWs is that, 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 Appendix Table A1, during our sample period 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), or 50% (for any irrevocable CLC that does not fall in the previous three categories). Of course, the fact that CCFs did not change during our sample period does not mean that their potential effect is nil: changes in CLC types that occur during the period that we study could affect bank capital (hence export flows) through CCFs. Ideally, we would like to control for such changes in the composition of CLCs. Unfortunately, the data at our disposal do not allow us to do so as we cannot observe the exact nature of export-related CLCs (i.e., whether they are confirmed-irrevocable, irrevocable-collateralized, non-binding, or another type). This is because the Turkish Ministry of Customs and Trade does not collect any information on the exact nature of CLC in the international trade transaction forms, which are the basis of the aggregate data made available to us by the TSI. Nevertheless, we believe that this potential shortcoming is not a major concern in our case for the following reasons. First, the overwhelming majority of CLCs used for Turkish exports are likely to be of either irrevocable-confirmed or irrevocable-collateralized types, both of which having less than one-year of maturity. This is because (i) non-binding CLCs have little use in international trade as they can be revoked when the exporter's bank requests payment, and (ii) international trade related CLCs are reported to have typically three to four months (but rarely above six-months) of maturity on average, i.e. much less than one-year of maturity (e.g., ICC, March 31, 2009 and October 26, 2011; and SWIFT October, 2009). Moreover, the Turkish regulation differentiates RWs applied to CLCs by *remaining* maturity (i.e., as of the date the bank starts holding them as an off-balance sheet position), rather than original maturity of the instrument at its issuance by the importer's bank. Moreover, the collateral feature only applies to certain manufactured goods (for ex., sheet iron and steel products) for which a commodity market exists. As a result, the overwhelming majority of CLCs in our data are either of irrevocable type for most industries (for which the CCF is equal to 100% during our sample period) or involve a manufactured good that can be collateralized (for which the CCF remains constant at 20%). This suggests that CLC types (hence CCFs) are industry specific: any CCF-related effects ought to be controlled by industry fixed-effects, which are a standard feature in all of our regressions. Consequently, our focus is in on Basel II-related changes in RWs.

Table 1 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 counterparty banks or to 100% for non-OECD-country based counterparty 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.[12] Unfortunately for us, the customs forms mandated by the Ministry of Customs and Trade, based on

[12] 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). For the sake of completeness, the corresponding benchmark-RWs, *irrespective* of the 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).

which the aggregated data are collected, do not collect counterparty foreign bank information. Instead, we rely on total asset (TA)-weighted average commercial bank ratings for each country.[13]

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 RW changes detailed in Table 1. Our hypothesis is that the banks would at least partially reflect these changes into CLC-clearing prices and/or ration (increase) CLC-confirmations, and exporters would react accordingly: we expect CLC-based exports to increase (decrease) when RWs decrease (increase).

First, suppose that a Turkish bank gets a request from a domestic corporate 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 sets aside $ 24,000 (approximately TL 43,400) in additional capital, *irrespective* of the risk of the counterparty-bank issuing the CLC (column A of Table 1).[14] After July 1, 2012, under Basel II, the same CLC's capital charge would depend on the OECD-based counterparty bank's agency rating (column D of Table 1). Suppose first that this agency rating is in the Aaa to Aa3 range (i.e., among top four investment grade categories) or non-rated.[15] 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 this group, which will form the base-case (control) group in our regressions. Now suppose that the counterparty bank rating is instead between

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[13] An alternative to TA-weighted bank ratings 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 a good proxy as the average of CLC-issuing banks' ratings in that country. Initially the Turkish banking regulator 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). However, during the implementation stage 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 (BDDK FAQ number 88). Nevertheless, we provide results based on sovereign ratings in Appendix Table A2.

[14] $ 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).

[15] While it might seem initially counterintuitive, the Basel Committee recommends applying the same RW for non-rated bank counterparties as the RW for investment-grade bank counterparties (see Table 1 for details). This is done in order to foster imports for low income countries that would otherwise be at a disadvantage due to the typically non-rated status of their financial institutions (BIS, October 2011). The Basel Committee also suggested the removal of the sovereign floor for the same reasons.

A1 and Baa3 (i.e., among the lower six investment grade categories). The corresponding RW would increase to 50%; hence the capital that the Turkish needs to set aside would go up to $ 60,000, a 150% increase.[16] In such cases, we would expect the cost of CLC-clearing increase for the bank, which would reflect this higher cost to its pricing for the service. This would potentially affect 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 CLC-requests for A1 and Baa3 rated bank counterparties more often because given the low margins involved in trade finance in general, holding export-CLCs become less attractive once the RW increases as if they were high risk products. In either case, there is no reason for OA- and CIA-financed exports to be affected by these RW changes.

We could also do a similar exercise when the same export-related confirmed-CLC is issued by a counterparty bank that is domiciled in a non-OECD country. Under Basel I, the capital that needs to be set aside by the Turkish bank was equal to $ 120,000 (column B of Table 1).[17] 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 (column D of Table 1). In the case of CLCs issued by Aaa to Aa3 rated counterparties located in non-OECD countries, we would expect exports to increase compared to the base case. If CLCs are issued by A1 to Baa3 rated or non-rated non-OECD banks, we would expect exports to increase, but less than that for Aaa to Aa3 counterparties given the smaller percentage decrease in capital charges for the lower investment-grade group. In contrast, exports involving CLCs issued by default-grade counterparties ought to decrease (but we have no such case in the sample, as described later). We also have a third identification scheme for CLCs to non-OECD countries if we assume that CLCs have, on average,

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[16] For OECD countries, there are two other possibilities in which the RW increases to 100% (below investment grade but non-default ratings) or to 150% (imminent or actual default). As mentioned earlier, these two cases are not covered here as they are eliminated from our sample given the very few countries that fall into those categories.

[17] $ 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).

remaining maturities less than three months (as depicted by the combination of columns B and C in Table 2).[18]

As a result, to test our two hypotheses we have three clear identification schemes (two of which apply to our non-OECD dataset, the third to our OECD data) 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 OECD or non-OECD country groups being differently affected depending on whether their banks are on average rated high-investment grade or low-investment grade) and depending on the average remaining maturities of CLCs.

We should note that changes to RWs applied to counterparty bank-issued CLCs need not necessarily affect export 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-corporate customers with repeated business with their institution were less likely to be affected compared to SMEs with less frequent export transactions. This is all the more so as some of the larger exporters, say in consumer durables, that belong to Turkish conglomerate groups, have their related-banks to request export financing from. Moreover, for a given bank, the RW changes need not necessarily lead to a material change in the amount of capital required for holding CLCs. This is because some of the RW changes (say, for A1 to Baa3 or non-rated OECD-domiciled bank counterparties) would require holding more capital, but others (say, for A1 to Baa3 or non-rated non-OECD-domiciled bank counterparties) would need less capital. The overall effect around Basel II

[18] In fact, 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 identification 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.

adoption would depend on Turkish banks' net exposure to CLCs in different rating ranges given the domiciliation (OECD or non-OECD) of the counterparty banks.

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.[19] Although, Basel II led to a decrease in the RBC ratios, the average effect was approximately an 1.5% drop, leaving the banking sector capitalized at roughly 15%. 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 much 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. Due to these data restrictions our tests are not as precise (hence subjected to higher standard errors) 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 RWs (hence capital requirements) per rating category differ depending on whether the CLC's *remaining* maturity is longer or shorter than 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 (ICC, 26 October, 2011, p. 16). So in our regressions we presume that majority of CLCs have maturities are either longer or shorter than three months, and run

[19] 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.

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.

Finally, we need to note that we cannot make welfare assessments regarding the impact of Basel II on total 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 (which excludes certain export-destination countries), and (ii) the identification schemes that differ for OECD and non-OECD countries (which require seperate regressions). However, our case would be strengthened if we were to observe results that are comparable across the OECD and non-OECD samples.

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 in Section 3.3), and (iii) some of the variables (such as import demand at the destination country excluding Turkish exports) that we use are not standard explanatory variables for gravity equations.[20] At any rate, our focus is on the difference-in-differences part of the empirical model.

[20] The data for GDP, which is the standard measure of destination import-demand, are available on an annual basis for most countries, but not at the semi-annual or quarterly level for all countries in our sample. 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 pre- and post-July 1, 2012.

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 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 *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 export-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 (all industries combined) of country c in during period t after excluding 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., $\exp(\alpha_3) - 1$).

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 case we obtain:

$$\begin{aligned} ln(EXPORTS_{c,i,t}) = {} & \alpha_0 + \alpha_1\, ln(IMPORTS\_EX\_TUR_{c,t}) + \alpha_2\, ln(DISTANCE_c) + \alpha_3\, D\_ADJACENT_c \\ & + \beta_1\, D\_A1\text{-}Baa3_c + \beta_2\, D\_BASELII_t + \beta_3\, D\_A1\text{-}Baa3_c \times D\_BASELII_t + \delta_i + \varepsilon_{c,i,t} \qquad (2) \end{aligned}$$

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;[21] $D\_BASELII_t$ 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 the average bank rating of the export-destination countries in our sample remain in 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 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 RW for OECD-country domiciled banks is 20% under *both* Basel I and II. It should be noted that the empirical model does not include (i) Ba1 to B3 (BB+ to B-), and (ii) Caa1 (CCC+) and below rating categories because there is no OECD country whose banks' average rating 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 (in terms of TA-weighted average bank rating as well as in terms of country's sovereign 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%). The non-rated category is also excluded from Eq. (2) because all OECD-member countries have ratings for at least some of their banks and in Bankscope we cannot observe non-rated banks issuing CLCs.

The coefficient estimates of interest are β_1 through β_3. Since these coefficient estimates 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

[21] 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.

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 that needs to be set aside 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 (either because banks reflect such changes to their CLC-clearing prices, or they simply ration such requests due to simple risk-control techniques based on rating classes).

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 inferences wrong, because we would be leaving 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 consider only CLC-based exports. A priori, in the absence of substitution between payment methods, 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 typically vary by country (as these differ in their demand for Turkish goods and shipment distance, hence shipment costs) and industry segment (due to Turkey's higher specialization in certain industries compared to others). 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.[22] Instead, Santos-Silva and Tenreyro (2006, 2011) suggest using Poisson or PPML estimators, 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.[23] 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:

[22] 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 this transformation leads to even higher biases in the log-linear gravity model estimates.

[23] 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).

$$EXPORTS_{c,i,t} = exp \{ \alpha_0 + \alpha_1 \, ln(IMPORTS\_EX\_TUR_{c,t}) + \alpha_2 \, ln(DISTANCE_c) + \alpha_3 \, D\_ADJACENT_c + \gamma_1 \, D\_A1\text{-}Baa3_c + \gamma_2 \, D\_BASELII_t + \gamma_3 \, D\_A1\text{-}Baa3_c \times D\_BASELII_t + [\beta_1 \, D\_A1\text{-}Baa3_c + \beta_2 \, D\_BASELII_t + \beta_3 \, D\_A1\text{-}Baa3_c \times D\_BASELII_t] \times D\_CLC_{c,i,t} + [\lambda_1 \, D\_A1\text{-}Baa3_c + \lambda_2 \, D\_BASELII_t + \lambda_3 \, D\_A1\text{-}Baa3_c \times D\_BASELII_t] \times D\_CIA_{c,i,t} + \delta_i \} + \varepsilon_{c,i,t} \quad (4)$$

where *exp{.}* denotes the exponential function and the rest of the variables are defined as above. 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 in Table 1 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, our hypothesis would suggest that 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 rated (A1 to Baa3 or non-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, which is a marginal destination for Turkish exports.

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:

$$
\begin{aligned}
EXPORTS_{c,i,t} = exp\{ & \alpha_0 + \alpha_1\, ln(IMPORTS\_EX\_TUR_{c,t}) + \alpha_2\, ln(DISTANCE_c) + \alpha_3\, D\_ADJACENT_c \\
& + \gamma_1\, D\_Aaa\text{-}Baa3\&NR_c + \gamma_2\, D\_BASELII_t + \gamma_3\, D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t \\
& + [\beta_1\, D\_Aaa\text{-}Baa3\&NR_c + \beta_2\, D\_BASELII_t + \beta_3\, D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t] \times D\_LC_{c,i,t} \\
& + [\lambda_1 D\_Aaa\text{-}Baa3\&NR_c + \lambda_2\, D\_BASELII_t + \lambda_3\, D\_Aaa\text{-}Baa3\&NR_c \times D\_BASELII_t] \times D\_CIA_{c,i,t} + \delta_i \} + \varepsilon_{c,i,t} \qquad (6)
\end{aligned}
$$

Finally, in other versions of equations (3), (4), (5) and (6), on top of industry fixed effects, we also introduce either (i) country fixed effects, (ii) country-time 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. TSI provided the exports data. These are based on (confidential) international trade transaction dataset that is maintained by the Turkish Ministry of Customs and Trade based on individual shipment documents that are filed electronically since 2006. We obtained quarterly exports data from TSI 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 aggregate quarterly country-industry-level exports data for 2011Q3-2012Q2 and 2012Q3-2013Q2 into pre- and post-Basel II annual 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 data 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 total assets weighted average bank ratings changed in such a way that they moved from one RW category into another (say, from AAA-Aa3 into A1-A3) some time between July 1, 2011 and June 30, 2013.[24] Second, we exclude Cuba, which retains its Caa1 long-term foreign currency sovereign debt rating from Moody's during our sample period, because it is (i) the only non-OECD 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, Syria and United Arab Emirates from our dataset. We drop Iran because, during the period of our study, Iran was subjected to an international embargo, which tilted this country's trade with Turkey in unusual ways towards gold (Financial Times, February 18, 2013). We also eliminate United Arab Emirates because most of the unusual Iranian gold transactions appear to have been done through this country (Financial Times, March 24, 2013). We exclude Syria, because that country's 2004 free trade agreement with Turkey was suspended on December 6, 2011 due to political differences over the handling of the Syrian civil unrest that turned into a full-blown civil war during our sample period. As a result, 2012 bilateral trade between the two countries shrank by 74% down to $ 566 million compared to its 2011 level. Fourth, we exclude two OECD countries (Hungary and Portugal) for which total assets weighted average bank ratings are below investment grade and another one (Greece) with a "junk" rating: we cannot draw meaningful inferences that can be generalized for these groups that contain so few countries. After all of these exclusions, the exports in our sample correspond to roughly 87.4% by value of those in the original dataset.

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[24] Barbados, Belize, Grenada, Luxembourg, Pakistan, Uruguay are excluded for this reason.

Long-term bank ratings by Moody's, S&P and Fitch are obtained from the Bankscope database.[25] 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.[26] For many banks we have ratings from more than one of these three agencies, 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 bank 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 is valid (in case there were changes to that institution's ratings over the annual period). 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 an alternative 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 total imports quarterly from all other countries after excluding shipments to country *c* 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 observation. As a result of these restrictions, for the OECD (non-OECD) sample we end up with exports to 29 (131) countries along 22 ISIC industry sectors under three different

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

[26] 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.

methods of payments (CIA, CLC and OA) for two years centered on July 1, 2012. The total number of non-OECD countries available for our time-varying country fixed effects regressions increases to 145 as we have 14 additional countries for which imports are not available in DOTS, yet these have TSI 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.[27]

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 the OECD (non-OECD) sample approximately 14.3% (37.5%) of country-industry-year observations in the dataset are equal to zero. The annual country-level total imports excluding shipments 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

[27] Note that *D_ADJACENT* never appears in Tables 3 through 6. This is because (i) in the regressions with the OECD sample reported in Tables 3 and 4, Greece, the only OECD-member-neighbor, drops out as it moves from default to speculative 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.

(the base-case rating-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 for non-OECD countries with unchanged RW). A similar distribution is obtained for the non-OECD group when we assume that the RWs for CLCs with maturities lower than three months.

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.[28] 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.[29] 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 of "basic metals". The example is consistent with the explanation provided by Antras and Foley (2014) for the small share of CLC-financed exports in the food industry: goods produced in the basic metal industry are easier to

[28] 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 (2014), examining export transactions of a large producer of frozen poultry products in the US, find that 6% of its exports are CLC-based.

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

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 a larger 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 type 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 rather than clustered standard errors at the country-level. This is because Kezdi (2004) suggests that approximately 50 clusters are needed for clustered standard errors to be efficient and our OECD sample has only 29 countries. In contrast, all of the non-OECD sample regressions are estimated with standard errors clustered at the country-level given the larger number of countries involved (131 or 145, depending on whether the specification includes $ln(IMPORTS\_EX\_TUR_{c,t})$ or time-varying country fixed-effects). In all of the PPML regressions we impose Stata's "strict" option, which prevents the problem of overfitting of the model to zero-trade observations. As a result, the number of observations may vary across columns for a given non-OECD sample depending on the set of fixed effects dropped from the empirical model by this Stata option.[30]

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 increases from 20% under Basel I to 50% (a 150% increase) for A1 to Baa3 rated (treated) counterparties under Basel II, compared to Aaa to Aa3

[30] It should be noted that only a limited number of zero-trade observations are dropped, so the problems associated with the presence of the zero industry-country exports still exists and requires PPML estimation.

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 in the specifications 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 used in international trade research. 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 only 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%. This finding is similar to the typical coefficient estimate of GDP in gravity models. In column A (C), the coefficient estimate for *ln(DISTANCE)* is equal to -0.923 (-0.967), both of which are statistically significant at the 1%-level. These coefficient estimates are also close to the estimates for distance observed in the trade literature. However, in column B (D), 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 (3.122) but not statistically significant. This is not too surprising as columns B and D look at *within* country variation of exports 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 columns A and C). 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.[31]

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* are equal to -4.302 and -4.339 in columns A and B, those for *D_CIA* are equal to -2.717 and -2.713 (all of which are statistically significant at the conventional levels): compared to OA-based exports, the incidence ratio for CLC-based exports is approximately -0.99 (= $e^{-4.3}$ - 1) and that for CIA-based exports is roughly -0.93 (= $e^{-2.7}$ - 1). These estimates confirme the much higher prevalence of OA-exports in the OECD exports sample. The coefficient estimates for *D_BASELII* are equal to 0.444 (marginally statistically significant at the 10%-level) and 0.467 (statistically significant at the 5%-level) in columns A and B: the incidence of OA-based shipments increase between 49% to 59% in the year after the adoption date on July 1, 2012. One possible explanation for this increase in the exporter (OA) financed exports is the reduction of RW for (domestic) loans to small to medium-sized enterprises (SMEs) under the new capital requirements. More specifically, SA version of Basel II stipulates that banks should decrease RW to SMEs with less than 50 employees and sales of less than 5 million TL ($ 2.766 million as of July 1, 2012) from 100% to 75%. This could generate a spillover effect from if the increase in SME lending would generate a larger availability of working capital, which in turn could lead to an increase exporter-financed OA shipments. We come back to this plausibility of such a scenario in the next section.

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 either: 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*

[31] These observations also hold true for the corresponding coefficients of equations (5) and (6) using the non-OECD sample (which are discussed under Section 4.2). To conserve space we only report the coefficient estimates for *D_CLC* and its interactions for equations (5) and (6). Full model estimates are available from the authors upon request.

is positive and statistically significant at the 1%-level: prior to Basel II adoption, there is more CLC-financed trade to OECD countries 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 coefficient estimates for the triple interaction *D_CLC×D_A1-Baa3 ×D_BASELII*, which is the focus of our test, are negative as we expected but not 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, tend to 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.

In columns C and D of Table 3, we present the results of PPML estimates with all of the possible 3,828 OECD observations. Our focus is now solely 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 in contrast to 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: it does not appear that lowering of the RWs for SMEs has had any impact on OA-based exports to OECD countries when the zero-trade observations are properly taken into account with a PPML estimator. These differences indicate that excluding zero-trade flows do indeed lead to biased estimates. We find that the coefficient estimates for *D_A1-Baa3×D_BASELII* in columns C and D changes sign compared to columns A and B but are 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, in columns C and D the coefficient estimates for *D_CLC* and *D_CIA* are approximately equal to -3.65 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% (= $e^{-3.65} - 1$) 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-Basel II 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 columns C and D the coefficient estimates are equal to -1.509 and 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%) as we hypothesized. We can calculate the associated RW elasticity of CLC-exports as -0.52 (= -0.78/1.50). This elasticity estimate suggests that a 1% increase in the RW for CLCs leads to roughly -0.52% drop in CLC-based exports to OECD countries with A1 to Baa3 rated banks on average. 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* remains, as expected, statistically insignificant in columns C through F of Table.

In Table 4 we examine the robustness of the Table 3 results for the OECD sample, by re-estimating Eq. (4) with different sub-samples. We do so by only reporting the results that correspond to the empirical model in the last column of Table 3 with separate time-varying industry and country fixed effects. 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.[32] 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 our estimation period). 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. These estimates are comforting in that the observed decrease in Table 3 for CLC-financed exports to OECD-member countries whose banks are rated A1 to Baa3 on average is driven by industry-country pairs that rely more on this type of payment (and not by some other unexplained feature of the data). 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. We do this exercise, with an admittedly ad hoc threshold of 95th percentile, to insure that our results are not driven by few outlier observations: the PPML estimator gives the same weight to all observations (Santos-Silva and Tenreyro, 2006), compared to, for example, the OLS estimator. In column C the coefficient estimate for *D_CLC×D_A1-Baa3×D_BASELII* is equal to -1.148, which suggests a -68% drop in the incidence of exports to destinations with A1 to Baa3 rated banks post-Basel II, which is in the same order of magnitude as the -79% incidence ratio of implied by column F of Table 3. We infer that our results are not due to some outlier export values during the post-Basel II period. 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 shipments by Turkish exporters to different destinations): looking at a sample with only the intensive margin (albeit only at the industry-country level and not at the firm level), we obtain very similar results compared to column F 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

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

-1.605 (statistically significant at the 5%-level). This finding suggests that the bias that might be due to autocorrelated errors (that Bertrand, Duflo, Mullanaithan, 2004, warn against in the case of the panel difference-in-differences models) does not appear to be important: we obtain a coefficient estimate that implies almost the same incidence ratio (-80%) as in column F of Table 3 (-79%). 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

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 country fixed effects) in a similar fashion to Table 4. 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. Given our hypothesis, 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% lower (= $e^{-1.316}$-1= -0.73) 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% lower

($=e^{-0.453}$-1), 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 non-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 countries 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. We come back on the economic interpretation of these elasticities (and those obtained for the OECD sample) under Section 5.

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 at the 10%-level), whereas the estimate of *D_CLC×D_A1-Baa3&NR×D_BASELII* is equal to 0.0440 but not statistically significant. We conclude that, in the case of non-OECD sample, the estimates for the triple interactions in column A are driven by industries that rely more on CLC-financing, as in the case of OECD sample. 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). The fact that the *D_CLC×D_Aaa-Aa3×D_BASELII* coefficient estimate is not statistically significant is not so surprising: for the non-OECD sample only 2.1% industry-country observations (including zeros) belong to the Aaa to Aa3 rating range (as opposed to 57.2% of the observations for the A1 to Baa3 and the non-rated categories). The coefficient estimate for the interaction *D_CLC× D_Aaa-Aa3×D_BASELII* is estimated with less precision in this case when compared with *D_CLC×D_A1-Baa3&NR×D_BASELII*. 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 these triple interactions.

In Table 6, we estimate Eq. (6) assuming that the remaining maturities for export-CLCs issued by non-OECD domiciled counterparties have maturities that 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, to conserve space 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 (and 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-Baa3 rated or non-rated non-OECD counterparty RW decreases from 100% to 20%, the incidence of exports to countries whose banks are investment grade or not-rated 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. Not surprisingly, the results are similar. 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 (but still statistically significant at the 5%-level) 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 0.52% increase in country-industry level exports to non-OECD countries whose banks remain investment grade or are not rated throughout our sample period. In column E, we reestimate Eq. (6) with a square panel, and observe that, as for the OECD sample, 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 we find no statistically significant effect as expected.

4.3 Regressions based on sovereign ratings

Finally, given a previous version of this paper, we re-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. The fact that the estimates typically corroborate our earlier findings but are weaker in terms of statistical significance is not surprising: sovereign debt ratings provide a poorer proxy for the ratings of a country's banks than the total assets-weighted average of bank ratings.

5. Interpretation of the results

In this section we evaluate the economic relevance of our results by discussing the possible channels that might explain our findings for the RW-change elasticity of CLC-based exports. These elasticities range approximately between -0.5 to -1 depending on the sample (OECD versus non-OECD) and the percentage changes in the RW moving Basel I to Basel II (which range from -50% to 150%). First, we should note that the estimated elasticities are specific to CLC-financed exports, and are not for the totality of Turkey's exports. Using a back of the envelope calculation and assuming that everything else can be held constant (for example, that there are no substitutions among different payment terms), we can estimate that the overall elasticity of total exports to RW changes to be between -0.032 and -0.179, given that CLC-export shares are on average 6.4% for the OECD countries and 17.9% for the non-OECD countries. 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 by Paravisini et al. (2014) for the reaction of Peru's total trade to financial shocks during the Great Recession.

One could make the counter-argument that the "shock" to the Turkish banks' capital through Basel II adoption in 2012 is much smaller in economic magnitude (given arguments made in Section 3.1 above) than the shock that the Peruvian banks suffered when foreign (mainly US) banks drastically decreased their supply of funding to foreign financial institutions. This is all the more so because Turkish banks were well capitalized around Basel II adoption. Our response to this criticism is three-fold.

First, a number of recent papers show that banks reflect the marginal cost of their capital to the pricing of their corporate loans.[33] For example, in line with the loan pricing theory by Repullo and Suarez (2004), Ruthenberg and Landskroner (2008) find that the adoption of Basel II in Israel rendered the pricing of corporate and retail loans more sensitive to the risk of the customer, especially for banks adopting IRB. In light of these results, it is not unreasonable to expect that banks would reflect the changes in the marginal cost of capital to the pricing of clearing and holding of CLCs presented to them by their corporate clients. In fact, Cosimano and Hakura (2011) use cross-country data and estimate that a one percentage point increase in equity to total assets ratio is associated with a 0.12% increase in average banking lending rates. Using a policy experiment in Brazil as an identification scheme, Martins and Schechtman (2014) find that an additional capital charge of 8.25% for a specific type of auto loans led to a 2.19% increase in related lending spreads (after using similar loans unaffected by the policy experiment as a control group). Martins and Schechtman (2014) also find that the auto loan rates were not reduced as much following the reversal of the policy experiment. Basten and Koch (2014) examine the effect of an additional one percentage point risk-weighted capital requirement put in place for mortgages on these loans' prices in Switzerland as part of Basel III adoption as of January 2013. They find that one percentage point additional risk-weighted capital requirement led to 0.17% increase in mortgage loan rates in capital constrained banks but had no

[33] It should be noted that these findings are in contrast with the basic premise of the Modigliani and Miller (1958) theorem on the irrelevance of financial policy for investment decisions in a frictionless world (with symmetric information, rational risk-based pricing of cash flows that are fixed and independent of capital structure, and no taxes – the usefulness of the paradigm being to study the economic effects of these frictions on capital structure and investment decisions; see for ex. Kashyap, Stein, and Hanson, 2010). The papers cited here suggest that the exogenous changes to capital structure affect pricing of loans even when other frictions that are known to exist remain constant.

effect on the prices of mortgages made by unconstrained banks.[34] It is clear that these results are not directly comparable with ours as they are obtained using different products (for ex., auto loans in Brazil for Martins and Schechtman, 2014; mortgages in Switzerland for Basten and Koch, 2014), with different maturities (two to five years for Brazilian auto loans, typically ten years for Swiss mortgages, versus approximately three months for Turkish export-CLCs), and risk, among other characteristics. Nevertheless, given that these findings suggest that regulatory changes in risk-weighted capital for a given type of loan are reflected in these loans' prices, it is not unreasonable to presume that the changes in RWs in the Turkish case would be reflected in the Turkish banks' pricing of export-CLC related services.

Second, in light of this empirical evidence, the question is how much of an elasticity should we expect given the change in the cost of CLC (including the change in the marginal cost of capital). To be able to answer to this question, we borrow from Schmidt-Eisenlohr (2013) an expression for the CLC-cost elasticity of exports:

$$\varepsilon = \frac{\partial EXPORTS}{\partial f^{CLC}} \times \frac{f^{CLC}}{EXPORTS} = (1-\sigma) \times \frac{f^{CLC}(1+r)^t}{(1+f^{CLC}(1+r)^t)} \tag{7}$$

where, f^{CLC} is the cost of CLC (as a percentage of transaction's value), σ is the (absolute) price elasticity of demand for exporter's manufactured goods, r is the foreign currency discount rate per year and t is the maturity in terms of years. Imbs and Méjean (2014) conduct a careful estimation of σ using industry-country level data and find its value to be between [6.64, 11.4].[35] There are few sources on the size of f^{CLC} as it is often a function of the risk ot the counterparty bank risk, and that of the

34 Some of the papers that rely on calibration exercises typically suggest smaller loan pricing effects following increases in capital requirements. For example, Kashyap, Stein, and Hanson (2010) conduct a calibration that suggests that a ten percentage point increase in capital requirements should lead to a 0.25% to 0.45% increase in lending rates. On the other hand, King (2010) conducts a different calibration that suggests that a one percentage point increase in capital ratio should lead to a 0.15% increase in loan spreads, a result that is comparable to Basten and Koch (2014), for example. That said, as Kashyap, Stein, and Hanson (2010) state, empirical studies based on shocks that are exogenous to lending (such as Ruthenberg and Landskroner, 2008; Martins and Schechtman, 2014; Basten and Koch, 2014) are likely to generate more reliable estimates of the impact of changes in capital structure on pricing and investment decisions, holding everything else constant. We note that a similar reasoning applies to CLC-based exports that we examine around Basel II adoption.

35 To the extent that Turkish producers have less pricing power and operate in more competitive segments, than say their peers from developed countries, the elasticity for Turkish exports should be on the higher end of the spectrum.

country of its domiciliation. One can find different quotes (from large international banks' websites) for import-CLCs that range from 1% to 8% of the transaction value with the costs in developed countries being at the lower bound and 8% being closer to values applied to developing countries or countries that face a financial crisis. It is not clear what would be the costs that the Turkish banks would charge to their exporter clients for clearing and holding non-Turkish counterparty bank CLCs.[36] We nevertheless limit ourselves to plausible but conservative values of f^{CLC} ranging from 1% to 5%. Finally, we assume an annual interest rate of 1.2% (in US dollar terms, as our exports data are in this currency), and a maturity of three months (i.e., t=0.25).[37] Based on these parameter values we obtain the following CLC-cost elasticity of exports from the numerical exercise: $\varepsilon(f^{CLC}=0.01, \sigma=6.64) = -0.06$; $\varepsilon(f^{CLC}=0.05, \sigma=6.64) = -0.27$; $\varepsilon(f^{CLC}=0.01, \sigma=11.4) = -0.10$; and $\varepsilon(f^{CLC}=0.05, \sigma=11.4) = -0.50$. Only the higher parameter values for f^{CLC} and σ get us a calibrated value for ε that is close to the lower estimate of ε that we obtain based on triple-interaction coefficient estimates and the percentage change in the RW (i.e., ε = -0.50). This suggests that the pricing channel explains part, but not all of the empirical CLC-cost elasticity of exports that we find.

This brings us to our third and last argument: the introduction or suppression of rationing. Basel II implementation requires that banks increase the risk-sensitivity of their equity capital. One way these institutions could adapt to the higher RBC requirements, given their existing level of equity capital, would be to implement risk measurement and control techniques, such as IRB, which can be fairly complex. In the meantime simple risk-control rules could be implemented before more sophisticated data-driven techniques (such as FIRB and AIRB) are built up internally. So it is quite plausible that following Basel II adoption in its SA version Turkish banks chose to ration (or remove rationing) holding export-CLCs at their corporate clients' requests if they deemed the counterparty banks to be more (less) risky, hence more (less) costly in terms of exposure. Such simple limits could

[36] The bankers that we contacted were reluctant to share this information, though one of them indicated that his institution requires a 10% profit margin over the actual transaction costs (e.g., wire transfer (SWIFT) fees, etc.) There is also heterogeneity among exporters: repeat customers of the bank and/or larger exporters are likely to face lower fees when compared with SMEs. Finally, we should note that while issuing a CLC to a Turkish importer involves stamp taxes and other legal fees (associated with the issuance of a financial guarantee), there are no such costs involved when holding a CLC that is already issued by a counterparty.

[37] The assumed interest rate has a highly marginal impact on the calculation: if r = 0%, the reported absolute elasticities do not change in a fall by less than 0.01.

be easily implemented by rationing (removing rationing) at the (i) counterparty (bank) level, (ii) at the sovereign country level, or (iii) at the rating class level.

In fact, a 2012 survey of the Asian Development Bank (detailed in Beck et al., 2013) suggests that rationing plays an important role in international trade finance decisions of banks. Survey repondent -- 106 large international Asian banks -- indicated that in 2011 they turned down 35% of the trade finance requests whose total value amounted to $ 4.6 billion. Survey participating banks also indicated the following obstacles as "significant" or "very significant" impediments to international trade finance (hence, arguably as a potential source for their rationing): (i) past-poor performance of the counterparty bank (83% of respondent banks), (ii) issuing bank's low credit rating (79%), (iii) low sovereign credit rating of the country where the issuing bank is domiciled (79%), (iv) Basel regulatory requirements (79%), and (v) issuing bank's low "capacity" (71%). In contrast, (i) only 29% of the respondents reported that the capital constraints that they face played a significant or very significant role in their decisions, whereas (ii) only 23% cited high transaction costs or low fee income as having a significant or very significant impact. While only suggestive, the 2012 survey results reported by Beck et al. (2013) indicate that rationing or its removal can play a very important role in banks' decision to accord trade finance. It could also be that some of the rationing occurs because banks do not find it profitable to provide CLC-clearing services given that SA version of Basel II treats short-term, low-risk CLCs as if they were long-term high-risk financial instruments. In fact, a recent ICC report (October 26, 2011) points out that for almost 389 thousand confirmed export CLC transactions in the survey sample (worth $ 195 billion) over 2008-2010, only 54 resulted in a default (which translates into 0.014% default rate) and only 19 of these resulted in an actual loss (equivalent to a loss rate of 0.035%). Combining this survey evidence with Kashyap, Stein, and Hanson (2010) observation that in competitive banking markets even a 20 basis points (i.e., 0.20%) difference in loan rates can lead to loss of corporate customers, we can make a reasonable case for this second potential rationale of rationing. For OECD-domiciled counterparty banks rated A1 to Baa3, when RWs increase from 20% to 50%, assuming that all confirmed CLCs have a CCF of 20% (the lower of the two possible values for our sample), there is a six percentage point increase in capital requirements [0.20×(0.50-0.20)]. This increase in cost may be enough to render holding export-CLCs on a bank's balance sheet non-profitable, even if these are still rated A1 through Baa3, i.e. investment grade. Similarly, assuming a CCF of 20% for non-OECD domiciled bank counterparties, a decrease in the RW from 100% to 50% (or to 20%) for Aaa-Aa3 (A1-Baa3) rated CLCs could amount to a 10% (16%) percentage point decrease in capital required for the position.[38] Such a large drop may lead the Turkish bank to stop its rationing and start accepting investment grade export-CLCs. Of course, this discussion does not exclude the possibility of a direct pricing channel being at work.

6. Conclusion

In this study 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 RBC requirements on banks' supply of loans has been documented, we know of no direct evidence of their impact on the real (manufacturing) sector besides Brun, Fraisse, and Thesmar (2014) and Lee and Stebunovs (2012). To the best of our knowledge, we are the first ones to show an economically relevant link between RBC requirements and the real economy via trade flows. Using industry-country level export 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 RWs associated with banks' holding of CLCs. In the Turkish case that we examine, the effect of Basel II adoption appears to be more subtle than that feared by instutitions involved in international trade. For example, exports to OECD countries whose banks are rated A1-Baa3 on average decrease following Basel II, whereas shipments to non-OECD countries whose banks are rated Aaa-Baa3 on average increase during the same period. Our evidence is robust to various specifications and estimations with various subsamples of the data and we find no impact with placebo tests. Importantly, we obtain similar elasticity estimates for CLC-based exports with two separate sets of countries (OECD and non-OECD) for which the identification schemes differ (with the related RW changes being in the

[38] It should be noted that we are not making a *direct* comparison of the 20 basis points suggested by Kashyap, Stein, and Hanson (2010) with six percentage change in capital requirement. Instead, we argue that the margin on CLC-clearing may become uninteresting given the capital requirements. In fact, a banker we spoke to said that his bank required that a 10% profit on export-CLCs on top of costs. Changes in RWs may render the clearing these products unattractive in terms of margin for certain institutions, which may prefer rationing instead.

opposite directions). Given recent evidence that loan prices are affected by capital requirements (e.g., Martins and Schechtman, 2014), our findings are consistent with a pricing channel for CLCs. A simple numerical exercise suggests that our results are unlikely to be explained by the pricing channel alone. Based on existing survey evidence, we argue that our findings are also compatible with a rationing channel, i.e., Turkish banks rationing CLC-acceptance when it becomes relatively more costly for them to do so. We provide a discussion on the plausibility of the pricing and rationing channels.

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, we can estimate that the cost of CLC-based exports would drop by 0.5% to 5% according to our elasticity estimates. 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 caused by the mechanisms we discussed above (which would not constitute the total change in imports induced by such a change). 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). These estimates imply a fall between 5.2% and 9.3% of total trade value (given the triple interaction estimates of columns D and A of Table 6, respectively) given the average CLC usage of 16.7% of total value of exports for the non-OECD countries whose banks are rated investment category on average (and assuming no substitution among the methods of payment).[39] 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.

[39] Here we can base ourselves on the triple-interaction (*D_CLC×D_Aaa-Baa3&NR ×D_BASELII*) coefficient estimate alone as the other related coefficient estimates for the control groups (say, OA after BASEL II adoption) are not statistically significant and close to zero whenever identified given the specification we use.

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 RBC regulation for banks. How to reconcile these findings 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 a "square" panel with all exports for a given industry-payment method-period set (6 observations) 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 1st, 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 method-period set (6 observations) 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 1st, 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 method-period set (6 observations) 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 1st, 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	