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## IMPACT OF DOMESTIC CREDITS ON THE CURRENT ACCOUNT BALANCE: A PANEL ARDL ANALYSIS FOR 15 OECD COUNTRIES

***A very fast credit growth that is denominated as a credit boom not only contributes to financial and economic growth, but also can have negative effects on the current account balance and can increase the probability of a financial crisis stemming from the current account deficit. This study aims to investigate whether domestic credits affect the current account balance. For this purpose, the panel data set of 15 OECD countries was used for the period 1986–2010. The results obtained from the panel ARDL analysis prove that domestic credits have a negative impact on the current account balance, as was expected.***

***Keywords:*** *credit boom, current account balance, panel ARDL analysis.*
***JEL Classification: C23; E51; F37.***

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## ВПЛИВ ВНУТРІШНЬОГО КРЕДИТУВАННЯ НА БАЛАНС РАХУНКІВ: ПАНЕЛЬНИЙ ARDL-АНАЛІЗ ПО 15 КРАЇНАХ ОЕСР

***У статті показано, що зростання внутрішнього кредитування, яке звичайно називають кредитним бумом, може не тільки сприяти фінансовому та загальноекономічному зростанню країни, але й мати негативні наслідки для балансу рахунків. Досліджено вплив внутрішнього кредитування на баланс рахунків за даними 15 країн ОЕСР у період з 1986 по 2010 рік. Результати панельного ARDL-аналізу доводять, що внутрішнє кредитування, як і очікувалось, має негативний вплив на баланс рахунків.***

***Ключові слова:*** *кредитний бум, баланс рахунків, панельний ARDL-аналіз.*
***Табл. 3. Форм. 7. Літ. 29.***

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## ВЛИЯНИЕ ВНУТРЕННЕГО КРЕДИТОВАНИЯ НА ТЕКУЩИЙ БАЛАНС СЧЕТОВ: ПАНЕЛЬНЫЙ ARDL-АНАЛИЗ ПО 15 СТРАНАМ ОЭСР

***В статье показано, что рост внутреннего кредитования, который обычно называют кредитным бумом, может не только способствовать финансовому и общеэкономическому росту страны, но и иметь негативные последствия для баланса счетов. Исследовано влияние внутреннего кредитования на баланс счетов по данным 15 стран ОЭСР в период с 1986 по 2010 год. Результаты панельного ARDL-анализа доказывают, что внутреннее кредитование, как и ожидалось, имеет негативное влияние на текущий баланс счетов.***

*Ключевые слова: кредитный бум, текущий баланс счетов, панельный ARDL-анализ.*

### 1. Introduction

If banks as financial institutions that collect funds from suppliers and lend them to the needed, lend to enterprises and to consumers at convenient conditions for both sides, investment and consumption expenditures will increase and economic activity will be fostered. Especially during stagnation or crisis periods, a decrease in bank credits may occur depending on increasing adverse selection and moral hazard prob-

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lems and so risk perceptions of banks, and this case is named a credit crunch. There are several studies in economic literature on credit phenomenon (Bernanke et al., 1991; Akhtar, 1993; Ghosh, Ghosh, 1999; Ikhide, 2003; Bulut, Erdem, 2013). These studies focus on the reasons of a credit crunch and investigate whether a decrease in credits indicate a credit crunch. Because a decrease in credits may also stem from the decrease in credit demand due to negative expectations of economic actors in the crisis environment.

Like a decrease in credits, a very fast growth of credits can also damage the economy, since a very fast credit expansion not only promotes investment and consumption and contributes to financial deepening, but can also result in balance sheet distortions and a financial crisis, the cost of which might exceed the benefits associated with the boom (Dell'Ariccia et al., 2012). A credit boom may lead to fragilities due to looser lending standards, excessive leverage, and asset price bubbles (Dell'Ariccia et al., 2012). In other words, while a rapid credit growth can deepen financial markets and stimulate the domestic demand, it can also increase the current account deficit, the fragility of the economy and the possibility of a financial crisis stemming from the current account deficit. A credit boom is identified with an abnormally high growth rate in private credits as % of GDP (Barajas et al., 2007). Because of the fact that financial development of a country is also measured by the credit-to-GDP ratio, and both Terrones (2004) and Dell'Ariccia et al. (2012) show a positive correlation between financial development and economic growth, it is important to detect whether the growth of this rate indicates a credit boom or not. Terrones (2004), Barajas et al. (2007), Mendoza and Terrones (2008), and Dell'Ariccia et al. (2012) examine this fact and present the cases which may be evaluated as credit booms. It is clear that a reasonable growth of credits is not enough or it does not assure financial stability automatically, but it can decrease the probability of a financial crisis relatively.

Because of the fact that the global financial crisis emerged in spite of price stability, and the very fast growth of mortgage credits in the US financial markets was one of the most important reasons of the crisis, different views concerning the main goal of central banks were suggested aftermath of the crisis. In this framework, Woodford (2011) argues that it is possible to generalize inflation targeting framework to take account of financial stability. Also, there are some other arguments that object to inflation targeting. For instance, Leijonhuvfud (2008) emphasizes that the monetary policy strategy in the US that was fundamentally only about controlling the price level failed as it led to the asset price bubble and a general deterioration in the quality of credit. Besides, Giavazzi and Giovannini (2010) argue that strict inflation targeting that ignores the financial fragility can produce interest rates that push the economy into a low-interest rate trap and can increase the likelihood of a financial crisis. Here what is significant is that writers remark the importance of financial stability.

This paper focuses on the relationship between the domestic credits and current account balances, and the purpose of this study is to investigate whether domestic credits affect the current account balances. The annual data from 1986 to 2010 are used as the period for the 15 OECD countries in the study. The rest of this paper is organized as follows: empirical literature on the relationship between credit booms

and financial crises will be presented in the first part. The second part is dedicated to the model and the data set. The methodology and the findings are covered in the third part. Finally, an assessment of the findings is presented in the conclusion part.

**2. Literature review**

A great number of empirical studies on the relation between credit booms and financial crises show that a credit boom increases financial instability, and it can even induce a financial crisis. Econometric methods used in these studies are surely important. When these methods are analyzed, it is seen that panel regression, panel logit, and panel probit models are mainly utilized.

Frankel and Rose (1996) applied the panel data set for over a hundred developing countries from 1971 to 1992 and found that rapid credit growth and overvaluation of real exchange rate were the two factors that induced currency crashes. Borio and Lowe (2002) utilized the annual data belonging to 16 developed countries for the period of 1960–1999. According to the results of their study, sustained rapid credit growth combined with large increases in asset prices enhances the probability of an episode of financial instability. Terrones (2004) examined credit booms in 28 emerging market economies from 1970 to 2002. The results indicate that credit booms are often associated with investment booms, current account deficits, and banking and currency crises and are typically followed by sharp economic downturns and financial crises. Kraft and Jankov (2005) examined the fast credit growth in Croatia in early 2000s and revealed that rapid credit growth increased the probability of credit quality deterioration and stimulated current account and foreign debt problems. Barajas et al. (2007) presented that larger and more prolonged booms and those coinciding with higher inflation rates and, to a lesser extent, low economic growth are more likely to end in crisis. The study covered the data belonging to 100 countries that experienced credit booms in different years. Mendoza and Terrones (2008) exploited the data belonging to 21 industrial countries that experienced 27 credit booms and 28 emerging market economies that experienced 22 credit booms for the period of 1960–2006. They defined the credit boom as an episode in which credits to the private sector grow more than during a typical business cycle expansion. The results show a systematic relation between credit booms and economic expansions, rising asset prices, real appreciations, widening external deficits and managed exchange rates. They also found out why all credit booms did not end in financial crises, why most emerging markets crises were associated with credit booms, and credit booms in emerging economies were often preceded by large capital inflows but not by financial reforms or productivity gains. Schularick and Taylor (2009) used a data set on 14 developed countries for the period of 1970–2008 and found that credit growth is a powerful predictor of a financial crisis and suggested that policy makers ignored credit at their peril. IMF (2010) examined the performance of emerging market economies during the 2008 global crisis. In the study, real domestic credit growth between 2003 and 2007 and credits to GDP in 2007 were used as two measures on pre-crisis overheating. According to the results of the analysis, countries that experienced pre-crisis credit booms experienced sharper output falls during the crisis, although to a lesser extent than during previous crisis episodes, and such credit booms were typically foreign-financed. Jorda et al. (2010) used almost the same data set which Schularick and Taylor (2009) did and came to a conclusion that credit growth is the best predictor of

financial instability, but the correlation between lending booms and current account imbalances has grown much tighter in recent decades. Dell'Ariccia et al. (2012) used a data set on 22 countries from 1998 to 2008 and found that while one-third of credit booms ended up in financial crises, others did not lead to busts but were followed by extended periods of below-trend economic growth. They also emphasized that bad booms tended to be larger and lasted longer even though it is difficult to differ a bad boom from a good one in real time.

**3. Model and data**

Based on the explanations above, the current account balance is described as a function of net domestic credits and an exchange rate. Accordingly, the empirical model is specified as follows:

$$CAB_{it} = \alpha_{0i} + \alpha_{1i} NDC_{it} + \alpha_{2i} lnEXC_{it} + u_{it}; \quad i=1, \ldots, N; t=1, \ldots T \qquad (1)$$

where $CAB_{it}$ is the current account balance as % of GDP, $NDC_{it}$ is net domestic credits as % of GDP, $lnEXC_{it}$ is the logarithmic form of a real effective exchange rate[4], and $u_{it}$ is the error term.

We use a panel data set on 15 countries (Australia, Austria, Canada, Finland, France, Germany, Israel, Japan, the Netherlands, Portugal, Spain, Sweden, Switzerland, the UK, and the USA) for the period of 1986–2010. The data are annual and obtained from the World Bank database. Expected signs of $\alpha_{1i}$ and $\alpha_{2i}$ are both negative.

**4. Methodology and findings**

Stationarity of the variables has to be examined before the panel data analysis and prior to time series analysis to avoid a spurious regression problem. In this context, Levin et al. (2002, LLC), Im et al. (2003, IPS), and Maddala and Wu (1999, Fisher-ADF) tests are widely utilized as panel unit root tests.

The LLC panel unit root test entails estimating the following panel model:

$$\Delta y_{it} = \mu_i + \rho y_{it-1} + \sum_{j=1}^{m} \alpha_j \Delta y_{it-j} + \delta_{it} + \theta_t + \varepsilon_{it} \qquad (2)$$

where $\Delta$ is the first difference operator, $m$ is the lag length, $\mu_i$ and $\theta_t$ are unit-specific fixed and time effects, respectively. The null hypothesis of $\rho = 0$ for all $i$ is tested against the alternative hypothesis of $\rho < 0$ for all $i$. The rejection of the null hypothesis indicates a panel stationary process.

$\rho$ is assumed to be homogenous for the LLC test. However, cross-sectional units can have a different speed of adjustment process towards the long-run equilibrium (Nazlioglu, Soytas, 2002). In this framework, Im et al. (2003) suggested a new panel unit root test that allows $\rho$ to vary across all. Therefore, the equation above is rewritten as follows while this test is being employed:

$$\Delta y_{it} = \mu_i + \rho_i y_{it-1} + \sum_{j=1}^{m} \alpha_j \Delta y_{it-j} + \delta_{it} + \theta_t + \varepsilon_{it} \qquad (3)$$

While the null hypothesis is $\rho = 0$ for all $i$, the alternative hypothesis is $\rho < 0$ for at least one $i$. The rejection of the null hypothesis indicates a panel stationary process.

[4] Nominal effective exchange rate is the value of a national currency with regard to a basket that consists of the currencies of the relevant country's trade partners. Real effective exhange rate is calculated by deflating nominal effective exhange rate. An increase in real effective exchange rate indicates the appreciation of the national currency.

Fisher ADF test proposed by Maddala and Wu (1999) combines the p-values from unit root tests for each cross section *i*. The test is non-parametric and has a chi-square distribution with *2n* degrees of freedom, where *n* is the number of countries in the panel. The test statistics is as below:

$$\lambda = -2\sum_{i=1}^{n} \log_e (\rho_i) \quad (4)$$

where $\rho_i$ is the p-value from the ADF unit root test for unit *i*. The Maddala and Wu (1999) test does not depend on different lag lengths in individual ADF regressions (Hossain, Saeki, 2012).

*Table 1.* **Results for Panel Unit Root Tests**

| Variable | LLC | | IPS | | Fisher-ADF | |
|---|---|---|---|---|---|---|
| | Constant | Constant and Trend | Constant | Constant and Trend | Constant | Constant and Trend |
| CAB | -0.147 (0.441) | 0.341 (0.6335) | -0.334 (0.368) | -1.065 (0.1433) | 30.799 (-0.4253) | 42.348 (-0.6668) |
| NDC | 2.892 (0.998) | 0.008 (0.5034) | 3.731 (0.999) | 2.202 (0.9862) | 17.600 (0.9647) | 14.557 (0.992) |
| lnEXC | -2.304 (0.010) | 0.148 (0.559) | -2.381 (0.008) | -1.120 (0.1312) | 45.184 (0.0371) | 36.392 (0.1955) |
| ΔCAB | -4.907 (0.000) | -2.944 (0.001) | -7.517 (0.000) | -5.107 (0.000) | 116.816 (0.000) | 81.969 (0.000) |
| ΔNDC | -2.264 (0.011) | -1.092 (0.1374) | -5.008 (0.000) | -4.513 (0.000) | 81.264 (0.000) | 72.939 (0.000) |
| ΔlnEXC | -8.943 (0.000) | -7.637 (0.000) | -7.772 (0.000) | -5.320 (0.000) | 116.522 (0.000) | 79.280 (0.000) |

*Notes:* Numbers in parentheses are prob. values. Δ is the first-difference operator. Newey-West Bandwidth selection with Bartlett kernel is used for both LLC tests. The SBC is used to determine the optimal lag lengths.
*Source:* Developed by the authors.

Table 1 shows the results of the LLC, IPS, and ADF Fisher panel unit root tests. Accordingly, while the current account balance and net domestic credits are stationary in their first differences, the results are not consistent for the real effective exchange rate. As known, in the event that all variables are stationary, fixed effects model or random effects model are estimated. Besides, panel FMOLS and panel DOLS must be employed if all variables are stationary in their first differences. Here, panel ARDL approach must be performed as this approach lets the estimation of the model in which the variables have different levels of stationarity.

Peseran et al. (1999) suggested two estimators for the ARDL model. The difference between these two estimators is that the mean group estimator (MGE) seems to be more consistent under the assumption that both slope and intercepts are allowed to vary across country, while pooled mean group estimator (PMGE) is consistent under the assumption of a long-run slope homogeneity (Ndambendia, Njoupouognigni, 2010). The basic assumptions of the PMG estimator are (Peseran et al., 1999): 1) the error terms are serially uncorrelated and are distributed independently of the regressors, 2) there is a long-run relationship between the dependent variable and explanatory variables, 3) the long-run parameters are the same across countries. This estimator is also flexible enough to allow for a long-run coefficient homogeneity over a single subset of regressor and/or countries (Simones, 2011). Peseran et

al. (1999) stated that the homogeneity test for long-term parameters can be performed by employing the Hausman test (1978). Under the long-term homogeneity assumption, even though both MG and PMG are consistent estimators, only PMG is the efficient estimator.

After these explanations, equation number 1 can be rewritten for panel ARDL ($p_i$, $q_i$, $k_i$) form as follows:

$$CAB_{it} = \alpha_i + \sum_{j=1}^{pi} \beta_{ij} CAB_{i,t-j} + \sum_{j=0}^{qi} \delta_{ij} NDC_{i,t-j} + \sum_{j=0}^{ki} \theta_{ij} \ln EXC_{i,t-j} + \varepsilon_{it} \quad (5)$$

As Peseran et al. (1999) argued, it is convenient to work with the reparameterization of the equation above:

$$\Delta CAB_{it} = \alpha_i + \beta_i CAB_{i,t-1} + \delta_i NDC_{it} + \theta_i \ln EXC_{it} + \sum_{j=1}^{pi} \beta_{ij}^{**} \Delta CAB_{i,t-j} + \\ + \sum_{j=0}^{qi} \delta_{ij}^{**} \Delta NDC_{i,t-j} + \sum_{j=0}^{ki} \theta_{ij}^{**} \Delta \ln EXC_{i,t-j} + \varepsilon_{it} \quad (6)$$

where $i = 1, \ldots, 15$, $t = 1986, \ldots, 2010$. $\varepsilon_{it}$ is the independent error term that was distributed on $i$ and $t$. $\delta^*_i$, $\theta^*_i$ and $\beta^{**}_{ij}$, $\delta^{**}_{ij}$, $\theta^{**}_{ij}$ represent long-term and short-term coefficients, respectively. Besides,

$$\beta_i = -\left(1 - \sum_{j=1}^{pi} \beta_{ij}\right), \delta_i = \sum_{j=0}^{qi} \delta_{ij}, \theta_i = \sum_{j=0}^{ki} \theta_{ij} \quad (7)$$

in the equation, $\varphi_i$ is the error correction coefficient and is expected to be significant and negative. The significant and negative coefficient of the error correction indicates that the error correction mechanism works.

*Table 2.* **Results for PMG and MG estimations**

| | PMG | MG | Hausman Test Statistics | Hausman Test p-val |
|---|---|---|---|---|
| Long-run Coefficients | | | | |
| NDC | -0.013** | -0.244*** | 0.97 | 0.33 |
| | (0.007) | (0.235) | | |
| lnEXC | -0.028** | 0.400 | 1.08 | 0.30 |
| | (0.014) | (0.411) | | |
| Error Correction Coefficent | | | | |
| $\varphi_i$ | -0.359* | -0.459* | | |
| Short-run Coefficients | | | | |
| ΔNDC | -0.049** | -0.040*** | | |
| | (0.030) | (0.032) | | |
| $\Delta NDC_{t-1}$ | -0.037** | -0.035** | | |
| | (0.025) | (0.022) | | |
| ΔlnEXC | 0.006 | 0.012 | | |
| | (0.021) | (0.031) | | |
| $\Delta \ln EXC_{t-1}$ | -0.006 | 0.007 | | |
| | (0.006) | (0.016) | | |
| Constant | -0.004 | | | |
| | (0.004) | | | |

*Notes:* The maximum number of lags for each variable is 3, and optimal lag lengths are selected by the AIC. The MG estimates are used as initial estimates of the long-run parameters for the pooled maximum likelihood estimation. The PMG estimators are computed by "back-substitution " algorithm. *,**, and *** indicate 1,5 and 10 % levels of significance, respectively. Standard errors are in parentheses.

*Source:* Developed by the authors.

Table 2 reports the results obtained from MG and PMG estimators. For the panel ARDL analysis, the Hausman test makes a choice between the PMG and MG estimators, and the null hypothesis of the Hausman test is that there is homogeneity. Because the null hypothesis can not be rejected, the PMG is the appropriate estimator for the model. Negative and statistically significant error correction coefficient ($\varphi_i$) shows the error correction mechanism works and indicates the presence of cointegration among variables. Accordingly, in case of instability, the dependent variable adjusts towards eqilibrium in almost 2.8 periods.

The negative and statistically significant coefficients of net domestic credits both in the short- and in the long-run indicate that net domestic credits affect the current account balance negatively as expected. Accordingly, it may be argued that the increase in net domestic credits create negative effects on the current account balance by stimulating the domestic demand and increasing expenditures. Of course, this may not mean a dangerous increase in current account deficits of the countries that form the group, but it found that an increase in domestic credits decreases the current account surpluses or increases the current account deficits of these countries.

*Table 3.* **Results of Diagnostic Tests**

| Country | PMG | | | | MG | | | |
|---|---|---|---|---|---|---|---|---|
| | $X^2_{SC}$ | $X^2_{HE}$ | $R^2$ | LL | $X^2_{SC}$ | $X^2_{HE}$ | $R^2$ | LL |
| Australia | 6.81 | 2.87 | 0.18 | 77.08 | 0.87 | 1.54 | 0.31 | 78.99 |
| Austria | 0.09 | 1.27 | -0.07 | 66.94 | 0.07 | 0.33 | 0.36 | 73.04 |
| Canada | 0.02 | 4.83 | 0.21 | 75.5 | 2.26 | 2.29 | 0.35 | 77.76 |
| Finland | 2.13 | 1.44 | 0.22 | 65.89 | 2.24 | 1.67 | 0.6 | 73.63 |
| France | 0.09 | 6.72 | -0.15 | 86.92 | 1.21 | 1.12 | 0.21 | 91.37 |
| Germany | 2.26 | 3.54 | 0.41 | 71.96 | 0.27 | 3.95 | 0.52 | 74.38 |
| Israel | 6.68 | 11.18 | 0.34 | 78.95 | 6.66 | 11.14 | 0.34 | 78.95 |
| Japan | 0.35 | 0.35 | 0.03 | 95.27 | 1.28 | 1.78 | 0.29 | 98.57 |
| Netherlands | 1.33 | 1.88 | 0.06 | 64.12 | 0.04 | 0.25 | 0.27 | 68.57 |
| Portugual | 1.67 | 0.01 | 0.41 | 62.73 | 1.13 | 0.28 | 0.15 | 64.28 |
| Spain | 14.12 | 0.77 | 0.24 | 63.26 | 0.68 | 5.42 | 0.65 | 72.17 |
| Sweden | 0.76 | 0.64 | 0.01 | 94.09 | 1.61 | 0.02 | 0.22 | 96.96 |
| Switzerland | 26.89 | 0.8 | 0.14 | 58.73 | 0.00 | 0.05 | 0.61 | 67.7 |
| UK | 0.16 | 0.09 | 0.28 | 66.38 | 0.05 | 0.1 | 0.29 | 66.48 |
| US | 2.11 | 0.02 | 0.14 | 83.27 | 0.14 | 2.43 | 0.43 | 88.33 |

*Notes:* $X^2_{SC}$: Breusch-Godfrey serial correlation test statistic, $X^2_{HE}$: White heteroscedasticity test statistics, $R^2$: Adjusted R-Squared, LL: log Likelihood.

*Source:* Developed by the authors.

In international trade theory, the J-curve effect is popular to examine the effect of a depreciation (or devaluation for a fixed exchange rate system) of domestic currency on the current account balance. Accordingly, because of the fact that export and import volumes do not change much in the short-run, the country receives less export revenue and spends more on imports, and this leads to a deterioration in the current account balance. However, export volumes start to increase and import volumes start to decrease, and as a result the current account deficit starts to improve and eventually moves into surplus in the long run (Pilbeam, 2006). While the negative effect in the short-run is denominated as a price effect, the positive impact in the medium- and long-runs is denominated as a volume effect. In this study, since we define the exchange rate as units of foreign currencies per unit of domestic currency,

we expect an inverse J-curve. Accordingly, on condition that the J-curve effect can explain the effects of exchange rate changes on the current account balance, a positive coefficient in the short-run and a negative coefficient in the long-run for the real effective exchange rate are expected in this study. However, the results show that while the real effective exchange rate does not affect the current account balance in the short-run, it affects negatively in long-run. Therefore, we can argue that there is no price effect, but there is a volume effect and the J-curve effect can not clarify the effects of changes in the exchange rate on the current account balance.

The results of the diagnostis tests in Table 3 show there is no autocorrelation or heteroscedasticity problems in the model.

**5. Conclusion**

In this study, we investigated whether net domestic credits and the real effective exchange rate affect the current account balance by using the annual data from 1986 to 2010 for 15 OECD countries. After the panel unit root tests, we utilized a panel ARDL approach suggested by Peseran et al. (1999) to obtain short-term and long-term coefficients.

The results of the panel ARDL approach indicates that net domestic credits affect the current account balance negatively in both short and long terms as expected because they stimulate the domestic demand and increase expenditures. When we examined the relationship between the current account balance and the real effective exchange rate, and revealed that real effective exchange rate affects the current account balance negatively in the long-run, and it does not have any effects in the short-run. As mentioned, these results do not support the J-curve effect because there are no short-term effects of the real effective exchange rate on the current account balance.

Finally, to the best of the authors' knowledge, this is the first study that analyzes the relation between the current account balance and net domestic credits by employing a modern econometric method such as the panel ARDL approach, for the studies in academic literature mainly focus on the relation between credit booms and financial crises. That is why, panel logit, panel probit and panel regression analyses are utilized in these studies. Therefore, we believe that this study might be useful for future empirical research on the relation between net domestic credits and the current account balance.

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Стаття надійшла до редакції 25.03.2013.