STOCKHOLM SCHOOL OF ECONOMICS

5350 Master Thesis in Economics

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Does owner-occupancy increase the unemployment rate?

A critical assessment of the Oswald effect

Abstract

The research community has been divided ever since Oswald (1996) first launched his finding of a positive relationship at the aggregate level between the unemployment rate and the share of owner-occupancy. Micro researchers question the finding altogether since numerous studies indicate that individual homeowners have better labor market outcome than tenants. This thesis questions the findings at the macro level and shows that the econometric techniques most commonly employed neglect the dynamics of the system as well as the endogeneity of the variables analyzed. The results from the OIRF analysis give support to the skeptics; the relationship is reversed in the short run and insignificantly positive in the long run. Hence, there does not seem to be a general contradiction for Swedish policy makers between promoting access to the housing market and advocating for reduced unemployment rates.

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JEL classification : C32, C33, J61, R23, R31
Keywords: owner-occupancy, unemployment, panel VARX, GMM

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Date of submission: January 7, 2014
Date of presentation: January 14, 2014

ACKNOWLEDGMENT

Firstly, I would like to thank my supervisor Rickard Sandberg, associate professor at the Stockholm School of Economics, for his valuable input during the course of writing this essay. I am especially grateful for his insightful comments regarding the methodology and for his general support, which have certainly helped to improve this thesis. Secondly, I also wish to thank Mikael Åsell, Martin Hill and Maria Vredin Johansson at the Ministry of Finance for their helpful comments throughout the process of writing this thesis as well as for providing lingual corrections.

The thesis has furthermore benefited from Inessa Love at the World Bank for providing her Stata codes to perform impulse-response and variance decomposition techniques in a panel VAR framework, from Timo Mitze, assistant professor at the University of Southern Denmark, for his input regarding the quantitative part of this thesis and from Monika Hjeds-Löfmark at Sieps for providing her input regarding the Swedish housing market.

Finally, I would like to thank my family and friends for their support. I am especially thankful to Meghan Geale who, despite the short notice, provided valuable lingual advice.

All remaining flaws and faults are my own responsibility.

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1. INTRODUCTION

The puzzling finding of a positive relation between homeownership and unemployment rates at the aggregate[1] level has been debated in the research community ever since Oswald first launched the finding in his paper 1996. Oswald argued that homeownership should not be favored at the expense of renting since this would have adverse effects by increasing unemployment rates. The increased transaction costs due to owning, induced by migration, he argued, would hamper the adjustment process after unemployment shocks. The Oswald effect, as it thereafter has been denoted, has attracted considerable critique from the research community mostly due to the findings at the micro[2] level of better employment outcomes of homeowners as compared to renters. However, at the macro level, results of previous research is mixed: some find a positive effect (Nickell 1998; Pehkonen 1999; Cochrane & Poot 2007; Isebaert, Heylen & Smolders 2013) while others discover insignificant results (Green & Hendershott 2001; Jonsson & Lindh 2012) or even dissident effects (Barrios-Garcia & Rodriguez-Hernandez 2004; Coulson & Fisher 2009; Lerbs 2010).

The question at hand is thus how the micro level findings of homeowners' superior labor market outcome as compared to tenants can coexist with macro level findings of a positive relationship between the unemployment rate and the share of homeownership. Five possible explanations have been suggested in the literature, arguing for the importance of the negative externalities induced by a large share of homeownership in a region.

Firstly, an increased share of homeownership might hamper the development of a well-functioning rental sector and might therefore decrease the flexibility in the labor market by increasing the moving costs for tenants as well as for the homeowners themselves (Oswald 1996; Isebaert, Heylen & Smolders 2013). Secondly, commuting over longer distances function as a substitute to moving, and even though at the individual level reservation wages increases with the commuting distance, commuting might still be associated with lower costs than moving due to homeownership. Several micro studies have confirmed that homeowners do in fact accept longer commutes (Kantor et al 2012). Hence, with a high share of homeownership, the individual commuting costs will increase due to congestion, as well as elevate the expenses of firms, which could undercut overall employment (Oswald 1996; Isebaert, Heylen & Smolders 2013). Thirdly, the increased interest of homeowners in the area surrounding the property might hamper efficiency in the labor market. Blanchflower and Oswald (2013) rest the hypothesis on the theory of "Not In My Backyard-ism" (NIMBYism), and state that homeowners might impede business activities that could lower property values (Fischel 2004). Fourthly, the lowered reservation wage and increased search intensity for local employment by homeowners due to mortgage payments might yield displacement effects (Blundel et al 2004; Crépon et al 2013; Laamanen 2013). The net effect on the unemployment rate would, however, depend on the amount of workers being displaced in relation to the number of homeowners finding employment. It also rests upon the assumption of less than perfect elastic labor demand and would in essence be a short-run effect (Laamanen 2013). Lastly, decreased consumption and increased savings due to debt-financed homeownership might constrain aggregate demand hence impeding the employment opportunities available (Laamanen 2013). In sum, these effects of homeownership could explain the seemingly puzzling finding of an Oswald effect observed at an aggregate level.

The discoveries of a positive relation would imply that the positive social effects of homeownership suggested by several authors[3] might come at the expense of decreased labor market opportunities at the aggregate level. A promotion of the housing market may thus augment the unemployment rate, an

[1] The level most commonly referred to as aggregate in the literature reviewed is regional levels such as municipalities and above, at the largest level it denotes the country level.
[2] The micro level is defined in the reviewed literature as the individual or household level.
[3] See among other Dietz and Haurin (2003) for a review of the social effects of homeownership.

effect that would have to be considered due to the severe social effects associated with unemployment as well as the adverse effects on long-term growth (Castells-Quintana & Royuela 2012). The political agenda of today in Sweden and across the developed world has in the aftermath of the most recent financial crisis largely been focusing on the persistently high unemployment rates. Any information of mechanism augmenting the regional unemployment rates would thus be of importance to policy makers.

The positive aggregate relationship of homeownership and unemployment rates is first and foremost an empirical phenomenon. It is not an extraction of a refined theory or economic model nor has it in full been implemented in an existing one[4]. Thus, its evidence is to a large extent dependent on the adequacy of the empirical techniques employed.

Recent empirics (see for instance Barrios-Garcia and Rodriguez-Hernandez 2004 and Holmes, Otero and Panagiotidis 2013) indicate that the macro level findings of a positive effect might be spurious resulting from omitted dynamics bias and an inability to properly control for endogeneity of the variables by instrumental techniques. Hence, the skeptics argue that the micro level findings are valid also at the macro level and that contrasting evidence is only the result of utilizing naïve econometrical techniques.

By the use of refined dynamic econometric models and sophisticated estimation techniques in comparison to the methods previously employed, the purpose of this thesis is to investigate whether there exists a positive relationship from homeownership to unemployment at the aggregate level utilizing Swedish regional NUTS-3[5] level data spanning from 1997-2011. The dynamic properties and endogeneity of the variables, a phenomenon recognized by several researchers but ignored by most empirical estimations of the linkages between housing and labor market variables, will thus be taken into account. Also, for comparability with previous literature, additional simpler methods will be utilized in the analysis along with the more refined. The research question is thus to examine if the phenomenon of an Oswald effect convincingly can be rendered as true and to explore the three explanations to this puzzle given by the research community: namely, reduced migration, increased labor supply and decreased amount of vacancies (the consumption theory and NIMBYism).

The main findings of this thesis are in accordance with, for instance, Barrios-Garcia and Rodriguez-Hernandez (2004) who employ a similar technique but with a restricted time length. From the analysis it is evident that the unemployment rate initially slumps after a positive shock to the owner-occupancy share. The unemployment rate thereafter returns to its pre-shock level after three years and then begins to increase. Neither the initial slump nor the subsequent increase is however significant at the 95 % level. The results are thus in apparent contrast to the results provided by among others Oswald (1996) and the hypothesis of an Oswald effect yielding a permanently increased natural rate of unemployment is thus rejected.

It could, however, be objected that one should be cautious in generalizing the results obtained in this thesis to the situation in other developed countries due to the regulated rental market in Sweden, see the *Discussion* section. Nevertheless, due to large similarities with previous international studies from the results of the simpler methods as well as Oswald's use of Swedish data in studies advocating for the effect, the reversed relationship and insignificant Oswald effect found in this thesis still provide

[4] The migration hypothesis has been implemented in the general equilibrium framework of Layard et al (1991) adding to the barriers of geographical mobility by Nickell (1998) and Nickell and Layard (1999). The empirical results by Nickell (1998) and Nickell and Layard (1999) indicate a positive relationship. However, the other aspects are not considered in their respective work.

[5] The NUTS (Nomenclature of Territorial Units for Statistics) is a European classification system of regional areas. The system consists of three levels and in Sweden the highest level, NUTS-1, consists of 3 regions ("Landsdelar") and the second most aggregate level is the NUTS-2 level consisting of 8 regions ("Riksområden"). The NUTS-3 level in Sweden consists of 21 regions referred to as "Län" in Swedish, see Table A3.

stark evidence for opposing the findings of econometrical models ignoring potential dynamics and endogeneity of the variables.
The remainder of the thesis is structured as follows. The next section reviews the relevant literature and the third section describes the hypotheses tested. In the fourth section the data is presented and the fifth section covers the empirical model utilized. The results are presented in the sixth section and the final remarks and conclusions are provided in the seventh and eight sections.

2. PRIOR RESEARCH

Prior to elaborating on the data used and the empirical estimation technique employed, a short review of the relevant literature for the simultaneous assessment of the housing and labor market is provided. A first point to stress is that the current literature does not rest upon an elaborated theory but rather tries to argue for potential explanations to an empirical dilemma by the use of empirically suggested mechanisms.

The literature within labor economics primarily rest upon the neoclassical framework of modeling an individual's lifetime expected utility maximization problem and explaining the dissipating of regional discrepancies in key labor market variables over time by labor mobility (Harris & Todaro 1972). The assumption underlying the theory of labor mobility as a clearing mechanism is one of rational individuals; a representative individual will only decide to migrate if by migration the individual can improve upon the current situation (see Etzo 2008 for a review of the literature).

In this setting, the inflow and outflow of labor in a region depends on the region's relative unemployment rate, relative real wage rate, moving costs and non-economic factors such as amenities. Hence, theoretically, geographical mobility should eliminate regional differences in labor market variables with higher wages and better employment opportunities encouraging an inflow of workers from less favorable labor markets (Blanchard & Katz 1992). Empirically, the latter factor has been determined the most important one (see SOU 2007:35 for a recent survey of Swedish studies), which might be explained by Treyz et al (1993) theory regarding risk-averse workers assigning more weight to the probability of employment than wage rates as a determinant of migration.

However, in the 1990s, economists struggled with the puzzling finding of seemingly persistent regional deviations from the national average unemployment rate. In theory, unemployment and wage disparities should disappear over time due to firm and labor mobility (Harris & Todaro 1972; Blanchard & Katz 1992). Oswald (1996, 1997a, 1997b, 1999) argued that the key to the finding of persistent regional labor market discrepancies was to be found in the housing market; he observed a sharp rise in the owner-occupancy rate in the U.S. and Europe during the same time span as labor market disparities had persistently prevailed. Oswald argued that the transaction costs involved in selling a home increase an owner's reservation wage and hence mitigates migration due to labor market disparities. The Oswald hypothesis, as it subsequently has been denoted, could thus contribute in explaining the sluggish adjustment to labor market shocks found at the regional level.

Oswald (1996, 1997, 1999) tested his hypothesis empirically by the use of OLS regressions, with and without controlling for fixed effect, investigating macro level data for 19 different countries spanning from 1969 to 1990. The overall results imply that an increased share of homeowners by ten percentage points increases unemployment by two percentage points. Oswald tested his hypothesis both on country and regional level and found, for the eight Swedish NUTS-2 regions, a significant relationship of the same magnitude as implied by the overall results.

The explanations to the observed pattern, Oswald argues, are severalfold. Ownership reduces mobility by increasing migratory transaction costs and thus negatively affects the geographical matching process in the labor market with increased unemployment rates as a result. Secondly, selling a home is associated with high transaction costs thus promoting commuting and, if commuting costs become high enough, unemployment may be a rationally better option. Thirdly, young people entering the labor market may become involuntary unemployed due to reduced mobility if the rental sector is small and homeownership is not within reach due to low financial opportunities. Lastly, Oswald argues, reduced mobility increases the possibility of skill-mismatch and hence lowers productivity, which might further affect the unemployment rate negatively.

Oswald's explanations to the paradoxical finding of a positive relationship between homeownership and unemployment at the aggregate level, hence rest upon both individual direct effects and indirect externalities as a high owner-occupancy share in a region affects the region as a whole and not only the homeowners.

The Oswald hypothesis has received substantial critique in the literature. Opponents argue that the estimation techniques employed lack in complexity resulting in heavily biased estimators or even spurious results (van Ommeren et al 2000; Dietz & Haurin 2003; BKN 2008). Furthermore, the findings at the micro level of better employment opportunities for homeowners as compared to renters are in apparent contrast with the direct individual effects in Oswald's theory[6].

In response to the critique, Oswald together with Blanchflower recently re-estimated his previous findings for the U.S. including various controls and controlling for potential hysteresis[7] in the unemployment rate and lagged effects of the owner-occupancy share on the unemployment rate. Blanchflower and Oswald (2013) once more find empirical support at the macro level for the Oswald effect; high levels of homeownership tend to increase unemployment rates with a one-year lag. Moreover, they expand upon the previously offered explanations for the connectivity, especially focusing on the externalities created by a high share of homeownership, and find that high homeownership rates are connected with lower labor mobility, longer commutes and fewer new firms.

Hence, the authors argue that the adverse effects on the unemployment rate of a high homeownership share function through reduced mobility of the labor force as a whole due to a reduced rental sector. Furthermore, the transaction costs, induced by moving, increase homeowners' rationale for commuting. Regions with high ownership rates will hence experience more commuting than otherwise comparable regions. Higher homeownership rates might thus increase regional unemployment rates due to increased reservation wages of all workers. The increased reservation wages would be due to, by Blanchflower and Oswald's line of reasoning, the increased traffic congestion negatively affecting all commuters. Moreover, by NIMBYism the authors further argue that a high degree of homeownership in an area might be associated with a reduced tolerance for new businesses thus increasing aggregate unemployment.

[6] See Coulson and Fisher (2002, 2009) for the US, van Leuvensteijn and Koning (2004) for the Netherlands, Munch et al. (2006, 2008) for Denmark, Jonsson and Lindh (2012) for Sweden and Laamanen (2013) for Finland.

[7] If the unemployment rate exhibits hysteresis, then its time path is characterized by non-stationarity. The literature is divided with regards to if the unemployment rate should be considered a stationary series in line with the natural rate of unemployment hypothesis or a non-stationary series following the hysteresis hypothesis. See Holmes Otero and Panagiotidis (2013) for a recent review of the literature.

In sum, Blanchflower and Oswald do not provide a throughout theory of the mechanism behind the observed relation between aggregate homeownership and unemployment but instead list possible explanations which might jointly contribute to the observed patterns.

Recently another study reached in essence the same conclusions with regards to high homeownership rates yielding higher unemployment rates. Laamanen (2013) study the deregulation of the Finnish rental market in the 1990s and finds that both previous micro and macro research could be correct. Laamanen argues that homeowners do have better labor market outcomes than renters but, ceteris paribus, regions with high homeownership rates tend to yield a causal positive effect on the overall regional unemployment rate. Hence a homeowner in a region with lower owner-occupancy shares has a more favorable labor market outcome than an identical homeowner in a region with a larger share of homeownership. Laamanen further tries to explain which mechanisms that might bring about this causal relationship from the regional homeownership share to the regional unemployment rate and finds empirical support for two hypotheses: externalities arising from consumption reductions and increased job competition by increased labor supply.

Laamanen rest the latter hypothesis on microeconomic findings and theories. Several researchers have argued that the favorable labor market outcomes of homeowners are due to their lower reservation wages and relatively active local job search for local employment arising from the need to meet high mortgage payments and reluctance to move to other regions (see for example Flatau et al 2003; Munch et al 2006). Recent research on displacement effects implies that this may cause negative externalities in the local labor market; as one group of individuals increases their labor supply other individuals may be displaced (Blundel et al 2004; Abbring & Heckman 2007; Crépon et al 2013). In sum, Laamanen argues that a higher homeownership rate leads to higher job search intensity and lower reservation wages which, under the assumption of less than perfect elastic labor demand, leads to externalities of a higher regional unemployment rate.

In support of the former hypothesis Laamanen argues that in addition to boosting homeowner's labor supply, mortgage loans may also affect consumption patterns. In recent studies of the effects of household credit on consumption behavior it has been shown that credit-constrained households with debt are forced to reduce their consumption when house prices are declining (Mian & Sufi 2010; Dynan 2012; Mian et al 2013). Laamanen argues that debt, under some circumstances, may be negatively linked to consumption even in the absence of unexpected changes in asset values. He refers to Stephens (2008) who show that repayment of a vehicle loan leads to an increase in nondurable consumption and to Coulibaly and Li (2006) who examine the effect of final mortgage payment and find that it is associated with an increase in durable consumption. Furthermore, Laamanen argues that mortgage loans may be a relatively profitable method of saving due to policies favoring homeownership in many countries. Owning has become the most common housing choice internationally and, as recognized by Atterhög (2005), this trend can by and large be explained by conscious politics; owning has been supported by subsidies, tax reliefs and interest rate subsidies. Hence, these reasons would illuminate why households do not want or are unable to save as much before they buy their home as after the purchase. Thus, it may be optimal for households to start saving more and consume less after a home purchase.

Moreover, a recent study by Mian and Sufi (2012) connects consumption decreases due to house price reductions to decreases in local employment. Hence, if an increase in homeownership and associated mortgage borrowing lead to a reduction in household consumption, this may be negatively reflected in the local labor market. Laamanen's consumption hypothesis thus is that higher homeownership leads to less spending which, through lower aggregate demand, leads to externalities in the form of an increased unemployment rate.

The Oswald effect has been subject to extensive research at both the micro and macro level. Table 1 provides a non-exhaustive list of the empirical research conducted utilizing macro level data. The overall results are mixed with regards to the existence of a positive relationship between homeownership and unemployment rates. Green and Hendershott (2001) and Jonsson and Lindh (2012) found a non-existent Oswald effect whereas Barrios-Garcia and Rodriguez-Hernandez (2004), Coulson and Fisher (2009), Lerbs (2010) and Holmes, Otero and Panagiotidis (2013) all found a negative relationship and Pehkonen (1999), Cochrane and Poot (2007) and Isebaert, Heylen and Smolders (2013) a positive one. Micro level research on the other hand almost unanimously rejects the existence of an Oswald effect at the individual level, see Munch et al. (2006) and Rouwendal and Nijkamp (2006) for recent reviews.

ARTICLES	Cross-section	Method	Effect
Oswald (1996, 1997a, 1997b, 1999)	Different countries and regions	Correlations, OLS and FE	Positive
Pehkonen (1999)	Finland	Bivariate OLS	Positive
		Multivariate OLS	No relation
Green & Hendershott (2001)	US	OLS by age group	No relation[a]
Flatau et al (2002)	Australia	Multivariate OLS	No relation
Barrios-Garcia & Rodriguez-Hernandez (2004)	Spain	3SLS	Negative
Cochrane & Poot (2007)	New Zealand	OLS, FE & Hausman Taylor	Positive
Coulson & Fisher (2009)	US	OLS & 2SLS	Negative
Lerbs (2011)	Germany	OLS	Negative
		FE	Positive
Jonsson & Lindh (2012)	Sweden	Correlation & OLS	No relation[b]
Blanchflower & Oswald (2013)	US	FE	Positive
Isebaert, Heylen & Smolders (2013)	Belgium	2SLS	Positive
Holmes, Otero & Panagiotidis (2013)	US	GIRF[c]	Negative
Horsewood & Dol (2013)	16 EU countries	GLS	No relation/Positive[d]

Table 1: Overview of studies of the Oswald effect

a. Positive significant relation found for the middle aged
b. Found an insignificant positive effect at higher aggregations but at municipality level the relation turns insignificantly negative
c. Analyzes the speed of adjustments to unemployment shocks by the use of generalized impulse-response functions (GIRFs)
d. Different results depending on included labor market variables

In sum, recent literature of Blanchflower and Oswald (2013) and Laamanen (2013) provide explanations for how the apparently contradicting findings at the micro and macro levels might coexist; an increasing degree of homeownership might create negative externalities that reverse the positive individual effects of homeownership. Blanchflower and Oswald (2013) acknowledge that higher ownership rates yield a slim rental sector that might lock-in workers, homeowners as well as tenants, in regions thus prohibiting migration and yielding a failure of the clearing mechanism of labor mobility. Moreover, Laamanen (2013) stress the need to consider the externalities at the aggregate level of an increased homeownership share by the consumption mechanism and displacement effect.

However, as recognized by Cochrane and Poot (2007), there was initially support at the macro level for the Oswald hypothesis (Oswald 1996; 1997; 1999; Pehkonen 1999) employing static correlation matrices and OLS regressions whereas more recent studies that expand upon the econometrical techniques have begun to question the findings altogether (Green & Hendershott 2001; Barrios-Garcia & Rodriguez-Hernandez 2004; Holmes, Otero & Panagiotidis 2013). The dynamic nature of the system and thus the endogeneity of the variables need to be accounted for in the empirical estimation, as stressed by Barrios-Garcia and Rodriguez-Hernandez (2004). Neglecting these features of the system might result in spurious results yielding "evidence" of an effect when there in fact is none. It could thus be that the Oswald effect is merely the result of omitted dynamics bias and the micro level evidence might thus hold also at the aggregate level.

Summarizing the empirical and theoretical research, there is no consensus in the current literature as regards the existence of a lock-in effect created by the housing market. However, the inability to properly control for the simultaneity of the variables and to take into account the dynamic effects is widespread in the empirical literature concerning the Oswald effect. It is therefore of interest to, in an econometrically sound way, assess the relationship between the housing and labor market.

3. HYPOTHESES

The purpose of this thesis is to assess the relationship between the housing and labor market by empirically investigating the validity of the Oswald effect. Prior research, reviewed in the previous section, generally rely on static panel estimations. Recent literature on the other hand has begun to question findings based on static estimations due to the likely endogeneity of the variables and the dynamic properties of the system of variables. Barrios-Garcia and Rodriguez-Hernandez (2004) is the only study that properly accounts for both the endogeneity of the unemployment rate and owner-occupancy share and the dynamic properties of the system[8]. It is hence of importance to assess the validity of an Oswald effect by considering the endogeneity and dynamic properties of the variables as well as to assess the three aforementioned explanations given by the literature for the relationship.

The research question is thus to examine if the positive aggregate relationship from the owner-occupancy share to the unemployment rate can be rendered as true and to explore the three different explanations provided by existing literature[9].

[8] However, Barrios-Garcia and Rodriguez-Hernandez (2004) only account for the dynamics in the unemployment equation and not in the homeownership equation. Furthermore, the study is restricted to one year, 1991, due to data restrictions on the ownership share. I however find it of importance to assess the effect over a longer time span due to the likely lags of the effect as well as restriction of generality induced by only assessing a specific year.

[9] Hypotheses 2-4 are in accordance with Blanchflower and Oswald's (2013) and Laamanen's (2013) hypotheses of the relationship between the labor and housing market.

Hypothesis 1
The Oswald effect, i.e. the null hypothesis is that a positive significant effect from the owner-occupation rate to the unemployment rate exists.

Hypothesis 2
The migration hypothesis, i.e. the null hypothesis is that a negative effect from the homeownership rate to the net migration rate exists.

Hypothesis 3
The displacement hypothesis, i.e. the null hypothesis is that a positive effect from the homeownership rate to the participation rate[10] *exists.*

Hypothesis 4
The consumption hypothesis, i.e. the null hypothesis is that a negative effect from the homeownership rate to the vacancy rate[11] *exists.*

These four hypotheses will be tested on Swedish data covering the period 1997 - 2011. The Oswald effect has by previous literature been tested in a large variety of developed countries and the Swedish NUTS-2 level regions are included in Oswald's empirical work from 1996. Furthermore, the effect has been found also with national level data as the aggregate. Hence, even though Oswald's latest work employ U.S. data, testing the Oswald effect using non-American data cannot a priori be considered non-adequate. However, the Swedish housing market is and has been regulated during the latest half-century, which might restrict comparability.

The Swedish housing market is characterized by a rather high degree of regulation, both in terms of price and quantity. Between 1959-1978 a "use value" system ("bruksvärdessystemet") was practiced, stipulating that the same rent would be charged for apartments with similar standard and size. This system generated rents for apartments in attractive locations far below the willingness to pay and the opposite for apartments in less attractive areas. During this period, the rental sector was also heavily subsidized (Jonsson & Lindh 2012). However, in subsequent years, neutralization has been adopted yielding a system that to some extent is pro-owner occupied housing (Jonsson & Lindh 2012).

Despite the potential limitations due to these characteristics, the Swedish housing market cannot be considered an invalid case for assessing the Oswald effect in line with international literature. Homeownership is characterized by high transaction costs and owner-occupancy is the most common housing form in Sweden in line with international tendencies. A further argument in favor of the appropriateness of using Swedish data is that Oswald (1996, 1997a, 1997b, 1999) employ Swedish data to argue for the existence of a positive effect.

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[10] The displacement effect could furthermore also be evident by an increase in the amount of hours worked. I do however not explore this possibility but instead focuses on the labor supply measured by the participation rate.

[11] The fourth hypothesis expresses the belief that the consumption effect can be captured by the job vacancy rate in a region; homeownership due to mortgage loans yields a reduction in consumption (in accordance with Laamanen 2013) which through decreased aggregate demand yields a higher unemployment rate by an reduction in the available vacancies. However, the aggregate demand effect on the vacancy shares might be counteracted by the fact that homeowners have high credit worthiness and thus might increase the business activity by starting up companies. Furthermore, in accordance with Oswald's hypothesis regarding NIMBYism (i.e. the argument that existing homeowners do not want businesses in their neighborhood that might adversely affect property values), the vacancy rate might be reduced by homeowners' negative attitude to new businesses and hence a significant finding of the vacancy rate might not solely capture the consumption mechanism.

To reach comparability with previous empirical assessments, the four stated hypotheses will be examined by the use of simpler static as well as more advanced dynamic models.

4. DATA

The annual data used in the empirical estimations was collected from Statistics Sweden and the Swedish Unemployment Agency. The regional level chosen for the analysis is NUTS-3 level and the time span was chosen to cover fifteen years, 1997-2011.[12]

The cross-sectional division of NUTS-3 level is considered to be well suited for the purpose of this thesis. Larger cross-sectional units, such as NUTS-2 level, might conceal the possible presence of an Oswald effect due to areas with low and high homeownership rates cancelling each other out. Smaller areas, such as municipalities, on the other hand, might not capture the externalities of a high owner-occupancy rate. The cross-sectional level chosen is thus adequate for the research question and there is no reason to believe that similar unit divisions such as labor market areas or functional units would alter the results to a significant degree[13]. However, it is beyond the scope of this thesis to conduct a full comparison of the difference in outcome due to the regional level chosen and the subsequent results should thus be interpreted acknowledging the cross-sectional division utilized.

The variables selected for the empirical model is further elaborated upon in the subsequent section about the chosen model. The section at hand expands upon the properties of the data. Table A1 provides a summary of the included variables and their sources and units whereas Table A2 contains the descriptive statistics of the variables. Most of the variables used are expressed in relation to the working age population, 20-64 years. Owner-occupancy is classified as including both homeowners and tenant-owners[14].

Stationarity of the time series is of importance in order to properly be able to interpret the subsequent findings from the empirical model. The presence of unit-roots in any of the series could generate spurious regressions or explosive results, indicating hysteresis and persistence in the response to innovations when there in fact is none. The usual conduct, when utilizing panel data, is to test for the presence of unit roots in any of the series using panel unit roots tests such as Im-Pesaran-Shin's IPS test (1997) and Levin-Lin-Chu's LLC test (2002). LLC assumes that all series are stationary under the alternative hypothesis while IPS assumes, under the alternative, that only a fraction of the series is stationary. The drawback of these tests is the strong null of unit-roots in all (LLC) or in a vast majority (IPS) of the regional series. The strong null thus creates a problem since rejecting the null might not give any indication of stationarity of the tested series (Enders 2010). Each of the regional series, for each variable, is thus tested separately using the Augmented Dickey Fuller (ADF) test. The ADF test has a null of non-stationarity and rejecting the null indicates a stationary series. However, due to time

[12] I wish to thank Tor Bengtsson at Statistics Sweden, Olle Westerlund at Umeå University and Kent Eliasson at Growth Analysis (Tillväxtanalys) for helpful guidance through the Swedish official statistical records.

[13] Branden (2008) also utilized NUTS-3 level data in her assessment of the Oswald effect at the micro level and Jonsson and Lindh (2012) find large similarities in correlations (and coefficients from OLS) between the unemployment and owner-occupancy share comparing NUTS-3 and LA level data.

[14] In Sweden, tenant-ownership ("bostadsrätt") is a form of owning, where the tenant-owner holds a share in a housing cooperative (a legal entity) that owns a real estate. It is often thought of as a form of homeownership since it may be viewed as a tradable asset, as recognized by Stephens (2003).

length limitations, the coefficient sign is primarily observed and strong conclusions are not made based upon the confidence intervals since the ADF test has a low power against stationary alternatives in small samples distorting the efficiency of the results (Mäki-Arvela 2003).

Migration and labor market variables are typically assumed to be stationary processes (Alecke, Mitze & Untiedt 2010). The total unemployment rate, participation rate, vacancy rate and net migration rate is thus expected to be stationary whereas the regional owner-occupancy series, the share of the population above 55 of age and the share of the population having undergone tertiary education do not yield any predetermined expectations of stationarity. The results are presented in Table A4 in the *Appendix* and are by and large in conformity with the a priori expectations. The unemployment rate, vacancy rate, participation rate, human capital share, migration and the variable of the older segment of the population are stationary series. However, the owner-occupancy shares do seem to have a problem of non-stationarity in one of the 21 regions. I, nevertheless, decide to continue with the series in their current form, thus not transforming the data further.

In summary, the majority of the variables are accepted as stationary processes by results from performed empirical test and a priori expectations. The next section will present the empirical model since stationarity of the series has been established.

5. EMPIRICAL MODEL

The empirical research on simultaneous assessments of key labor market variables can be divided into two main approaches; structural (see Mitze and Stephan 2013 for an overview) or time-series techniques (Blanchard & Katz 1992; Decressin & Fatas 1995; Mäki-Arvela 2003; Barcellos 2009; Boubtane, Coulibaly & Rault 2012). A structural approach involves the ex-ante classification of variables as being either endogenous or exogenous. This approach, relying on a rather ad-hoc categorization of variables employed to instrument the contemporaneous endogenous explanatory variables, has gained substantial critique recently (Rickman 2010; Mitze & Stephan 2013). The alternative approach, exploiting the time-series properties of the data set, is thus often considered superior since it, in its unrestricted form, treats all included variables as endogenous (Mitze & Stephan 2013). A further advantage of the vector autoregressive (VAR) technique is the use of impulse-response analysis to explore the dynamic properties of the system.

The panel VAR[15] method exploits both the cross-sectional and time-series nature of the panel data set which is, as argued by Holtz-Eakin et al (1988), perfectly fitted for VARs due to the reduced requirement of length for the time dimension. Binder et al (2005) provide a throughout discussion regarding various estimators for panel VAR models in a small *T* large *N* setting, and Canova and Ciccarelli (2013) provide an updated survey of the different panel VAR techniques employed for the analysis of heterogeneous units.

Utilizing a dynamic panel model to explore the structure of the labor market with respect to the housing market would supplement the traditional analysis, mostly employing static models. As argued by Bond (2002) and Baum (2006), static panel estimates and OLS models omit the dynamic properties

[15] Regarding the notation, a panel vector autoregressive model is usually denoted panel VAR or PVAR and is a model utilizing solely endogenous variables. The VARX model allows for the inclusion of exogenous variables in the model. See appendix A5 for a throughout discussion.

of the system, which could cause a problem of dynamic panel bias, and also prohibit any analyze of the adjustment process (Baltagi 2008). Omitted dynamics is a serious misspecification of the model which could result in spurious results (Bond 2002). Furthermore, a setting with N larger than T is ideally suited for panel dynamic modeling as argued by among others Bond (2002), Baum (2006) and Roodman (2009). Lastly, a potential problem of endogeneity is easier to control for in a dynamic panel model compared to a static model and OLS that do not allow for internally generated instruments (Greene 2008). The dynamic panel model could hence be argued as preferable to static panel modeling and standard OLS for the research question at hand, since the dynamic panel model accounts for additional features of the series that if not controlled for might yield false interference.

A generalization of the econometric model employed in the analysis has the following reduced form

$$Y_{it} = A_{i0} + \mathrm{A}_{i1} Y_{i,t-1} + \mathrm{B}_{i,1} X_{i,t-1} + \varepsilon_{it} \quad (1)$$

where $i \in \{1, 2, \ldots, N\}$, $t \in \{1, 2, \ldots, T\}$, Y_{it} is a vector of stationary endogenous variables for M equations, A_I and B_I are matrix polynomials in the lag operator and contain the respective coefficients, X_{it} is a vector of stationary exogenous variables for M equations, A_{i0} encompasses constant effects and ε_{it} is a vector of white noise error terms.

The research question specified is to analyze the relationship between the housing and labor market and the empirical model thus includes the share of owner-occupied housing (*own*), total unemployment rate (*utot*), net migration (*net*), participation rate (*part*) and vacancy rate (*vac*) in line with Blanchflower and Oswald's (2013) and Laamanen's (2013) research as well as previous empirical assessments of the Oswald effect (see for instance Barrios-Garcia and Rodriguez-Hernandez 2004). These variables are expected to depend on specific regional characteristics such as the demographical composition. Generally speaking, homeownership is concentrated to the higher age segment of the population while the younger part of the population typically is seen as less employable than their older peers (Rosen 1979; Heiborn 1998; Green & Henderscott 2001; Barrios-Garcia & Rodriguez-Hernandez 2004). To properly account for the demographical setting in a region when estimating the Oswald effect, the share of the population above 55 years of age in relation to the population as a whole (*pop55*) is included in the dynamic model.

Moreover, the skill level of the regional population is likely to affect both the owner-occupancy share as well as the labor market variables. As Dohmen (2005) emphasizes, interregional migration of workers only exists when the wage exceeds the unemployment benefits and cost of reallocating into the migration regions compared to the current region of residence. Hence, since wages are higher for highly skilled workers the benefit of reallocating may thus by far exceed the costs for this sub-segment of the population. The higher degree of mobility among the highly educated has been confirmed for the Swedish population by among others the Nordic Council of Ministers (2002). Dohmen's (2005) theoretical model thus implies that it is crucial to control for the skill-level of the population when empirically assessing the relationship between the labor and housing market. The share of tertiary educated in the region in relation to the working age population as a whole (*hc*) is thus included in the empirical model in line with among others Cochrane and Poot (2007) and Isebaert, Heylen and Smolders (2013).

The inclusion of demographical factors and skill-composition of the regional population is in line with prior research of the Oswald effect by for instance Barrios-Garcia and Rodriguez-Hernandez (2004) and Isebaert, Heylen and Smolders (2013). The reason for why control variables such as the regional demographic and skill composition are consider vital for the validity of the model is the likely problem of omitted variables bias (OVB). The effect, if any, of the control variables on the left-hand side variables might be taken into account by the coefficient of the right-hand side variables if the controls are omitted, thus creating a problem of correlation between the right-hand side variables and the error term. Correlation between the error term and the right-hand side variables is a misspecification of the model, and the regional skill level and demographical composition is included in the model to control for a potential problem of OVB.

The system of equations below provides the expression for the PVARX (1) model used to empirically assess the research question of this thesis.

$$\begin{gathered} Y_{it} = A_0 + A_1 Y_{i,t-1} + B_1 X_{i,t-1} + \varepsilon_t \qquad (2) \\ Y_{it} = [own_{it},\ utot_{it},\ net_{it},\ part_{it},\ vac_{it}]' \\ X_{it} = [pop55_{it},\ hc_{it}]' \\ A_0 = [year, region] \\ i \in \{1,2,\dots,21\} \\ t \in \{1,2,\dots,15\} \end{gathered}$$

The above specification is largely consistent with previous research (Cochrane & Poot 2007; Isebeart, Heylen & Smolders 2013) but with the inclusion of time operators in line with Barrios-Garcia and Rodriguez-Hernandez (2004)[16]. The static equations, with the labor market variables expressed as purely independent variables, are frequently used in empirically estimations of the Oswald effect. An analysis depending solely on static equations might be misleading by the joint determination of the variables, as further elaborated upon throughout this section, yielding potentially biased estimates due to omitted dynamics (Bond 2002). An empirical scheme, including an extensive model with variables chosen based on previous literature and an assessment of the dynamic relationships without the need for ex-ante restrictions on signs and causality direction, has thus been chosen for the purpose of this thesis.

The hypothesized signs of the coefficients are coherent with the hypotheses outlined in the third section. The main purpose is to explore if the Oswald effect can be considered a non-spurious effect in the Swedish data analyzed. The hypothesized effect is thus that an increase in the lagged share of homeownership should have a positive impact on the unemployment rate. Furthermore, an increase in the lagged share of owner-occupancy should negatively influence the net migration rate and vacancy share, and positively affect the participation rate.

Analyzing the magnitude, sign and significance of the coefficients included in the PVARX(1) model serve their purposes and provide guidance on the importance of the relations between the labor and housing market. Nevertheless, the main contribution and the important distinction with regard to the existing literature is to examine the dynamic relationships through impulse-response graphs, tracking the effect of shocks to the owner-occupancy share on the labor market variables.

[16] Other valuable variables could however have been added to the empirical model such as the real wage rate, user costs, unionization, the share of households with children, share of widows and singles, other age compositions, GRP, population density etc. However, additionally including these would have resulted in a severe drop in degrees of freedom which might erode the possibility of significant results even though the true point estimate might be significant. Accounting for the potential problem of an excessive amount of included variables I chose to utilize a model including the variables that are, by my opinion, the most relevant for the question at hand.

A straightforward way to explore the magnitude of different shocks to the system is to graph the impulse-response functions (IRFs) of the system. The IRFs is a function of the coefficients and traces the typical reaction of the system to a specific unit shock of a variable's error term in the system. However, interpreting the IRF graphs is usually difficult due to the likely correlation between the equations error terms. For example, in examining the likely response of the regional unemployment rate due to a unit shock to the homeownership share's error term, everything else in the system cannot convincingly be regarded as constant due to the contemporaneous correlation between the error terms in the system. One way to overcome this problem is to assume a casual ordering of the variables in the system by implementing recursive assumptions to construct mutually uncorrelated innovations. Imposing these restrictions enable the different equations' residuals to be uncorrelated, i.e. orthogonal, yielding orthogonalized impulse response functions, OIRFs[17] (Hamilton 1994). This is particularly important for the purpose of this thesis since correlated errors are allowed for a priori.

Following the process outlined in Love and Zicchino (2006) and Alecke, Mitze and Untiedt (2010), the OIRFs can be derived in the following way. The moving average representation of the panel VARX system is expressed as

$$y_t = \xi + \sum_{i=0}^{\infty} \Gamma_1^i \varepsilon_{t-i} \qquad (3)$$

where ξ is a function of model parameters[18]. Due to co-movement in the error terms, the expression can be rewritten as

$$y_t = \xi + \sum_{i=0}^{\infty} \kappa_i \epsilon_{t-i} \qquad (4)$$

where κ_i is the impulse response coefficients. The elements of κ_i thus express the impact of one unit shock in a specific error term. To quantify the cumulative response of an uncorrelated shock, the error terms need to be orthogonalized. Continuing the Love and Zicchino (2006) and Alecke, Mitze and Untiedt (2010) approach, it is assumed that $\Omega = E(\epsilon_t \epsilon_t')$ is positive definite and thus define $u_t = K^{-1}\epsilon_t$ with K being a lower triangular matrix with ones along the principle diagonal. The expression materializes into

$$y_t = \xi + \sum_{i=0}^{\infty} K\kappa_i \epsilon_{t-i} \qquad (5)$$

For any shock in $\epsilon_{x,t-s}$ of variable x, the response of variable y can then be written as

$$\frac{\delta y_t}{\delta \epsilon_{x,t-s}} = \kappa_s K_x \qquad (6)$$

[17] I am grateful to Inessa Love for providing her Stata code for the OIRF graphs and variance-decomposition (Love & Zicchino 2006).

[18] The exogenous variables are also incorporated in ξ, however since this thesis' main concern is the impulse and response of the endogenous variables they are not explicitly modeled for.

where K_x represents the first column of matrix K, defined above. It is important to note that the orthogonalization of the residuals demands an explicitly determined ordering by which any correlation between residuals of any two elements of the variables is allocated to the first appearing variable in the ordering.

Hence, by employing the Cholesky decomposition to orthogonalize the error terms, the OIRFs are calculated by tracing a one standard deviation sized innovation through the equation system. As mentioned, a shortcoming with the procedure is that the ordering of the variables determines the degree of endogeneity of the different variables. The ordering is thus important for the results obtained and cannot be determined via statistical procedures invoking uncertainty into the model. I regard the housing variable as more exogenous than the labor market variables and thus choose the following ordering $[own_{it} \rightarrow utot_{it} \rightarrow vac_{it} \rightarrow net_{it} \rightarrow part_{it}]$. [19]

In addition, the variance-decomposition based on the OIRFs will also be reported. The variance-decomposition yields the proportion of movement in one variable being due to shocks in the own variable and the rest of the system variables. The decomposition shows to what degree an s-step ahead forecast error variance for each variable in the system is explained by the innovations to each variable. In sum, the OIRFs provide information concerning the reaction to innovations by the system, i.e. the dynamic properties, whereas the variance-decomposition displays the relative importance of each innovation to the variables in the system.

Before turning to the analysis of the obtained results, a short description of the chosen estimator and the instrumental technique employed will be provided in the subsequent paragraphs.

The generalized method of moments (GMM) technique was developed by Lars Peter Hansen in 1982 and is a generic estimation procedure. The GMM estimators are consistent, asymptotically normal and efficient under standard regularity conditions. The method necessitates that a certain number of moment conditions are specified, which are functions of the model parameters and the data such that their expected value is zero at the true parameter values. The GMM method thus minimizes a certain norm of the sample average of the moment conditions[20] (Hansen 1982).

There are several advantages with GMM techniques, in comparison to alternative cross-sectional estimation techniques. The GMM estimator accounts for the dynamic structure of the process in which current realizations of the dependent variable may be influenced by past ones. Furthermore, arbitrarily distributed fixed individual effects may be a concern in standard cross-sectional regressions, which essentially must assume these effects away. In a panel setup, the time variation can be used to identify the parameters (Roodman 2009). Moreover, the estimator allows for both endogenous and strictly exogenous regressors. Finally, the idiosyncratic disturbances are allowed to have individual-specific patterns of heteroskedasticity and serial correlation but are assumed to be uncorrelated across individuals (Roodman 2009). The GMM estimator is thus suitable for the purpose of this thesis even though the assumption of uncorrelated cross-sections might be too strong given the data and question at hand.

[19] I however altered the ordering of the labor market variables as a robustness control and the main implication of the OIRF analysis was then not significantly changed, results to be provided upon request.

[20] The properties underlying the GMM technique are elaborated upon in more detail in the technical appendix, A5.

The identification of the GMM estimator relies crucially on the instrument matrix applied for the endogenous regressors. Arellano and Bond (1991) propose a GMM estimator that makes use of all lagged variables for the instrument matrix resulting in the following "standard moment conditions":

$$E\left(Y_{i,t-p}\Delta\varepsilon_{it}\right) = 0 \ \forall \ p = 2, \dots, t-1 \tag{7}$$

Where Y is the stationary endogenous variables, t denotes time, p the lag length of the utilized instruments and ε is the error terms. However, as Blundell and Bond (1998) argue, GMM estimators in first differences, generally denoted DGMM, might behave poorly since lagged levels could provide weak instruments for subsequent first differences. Blundell and Bond thus develop the "stationarity moment conditions", which uses instruments in first differences for the equation in levels:

$$E\left(\Delta Y_{i,t-1}\varepsilon_{it}\right) = 0 \ \forall \ t = 3, \dots, T \tag{8}$$

Blundell and Bond (1998) suggest the use of a system GMM estimator (SGMM) that jointly includes the standard and stationarity moment conditions. The instrument set for the case of SGMM is thus defined as $Z_i = \left(Z_i^{\Delta}, Z_i^{L}\right)$, where the first term Z_i^{Δ} denotes the instruments for the variables in first differences in line with equation (7) and the second term Z_i^{L} includes the first differenced instruments for the variables in levels as expressed in equation (8).

In the present context the SGMM is preferable to the DGMM estimator. The main advantage of SGMM, as noted by Baltagi (2008), is that it generally produces more efficient and precise estimates by reducing the finite sample bias. Moreover, the SGMM method is superior to the DGMM when the included variables are random-walks or close to random walks (Bond 2002; Baum 2006; Roodman 2009). Since the ADF tests of stationarity revealed a potential problem in the housing variable, the SGMM estimator is the more appropriate of the two. It is however important to note that the stationarity moment conditions depend on the first moments of the time series being time-invariant conditional on common year dummies. The chosen estimation period, 1997-2011, nevertheless seem to correspond well with the assumption due to fairly stable regional disparities in the variables at the beginning of the period. Furthermore, the included year dummies are expected to control for any problem of instability and their inclusion is thus expected to mitigate any potential problem in this respect.

The GMM estimator can furthermore be estimated by the use of single equation generalized method of moments (SE-GMM) or multiple equation generalized method of moments (ME-GMM) technique with the former being a special case of the more general ME-GMM technique. The SE-GMM estimates the system of equations in an equation-by-equation framework whereas the ME-GMM technique stacks the equations into a system thus accounting for correlation between equations (Hayashi 2000).

In the estimation of the dynamic model, expressed in equation (2), ME-GMM is employed by stacking the system of equations by the use of three-stage least square (3SLS) estimation and instrumenting the regressors by SGMM technique[21]. The instrument matrix is defined as below for a system of *M* equations.

$$Z_i^S = \begin{bmatrix} Z_{i1} & \cdots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \cdots & Z_{iM} \end{bmatrix}$$

Where $Z_{im} = \left(Z_{im}^{\Delta}, Z_{im}^{L}\right)$ for $m \in \{1,\ 2, \dots, M\}$. Z_{im}^{Δ} contain the lagged variables, previously denoted *Y* and *X*, to instrument the endogenous variables, *Y*, in first differences whereas Z_{im}^{L} denotes the lagged first differenced variables, *Y* and *X*, to instrument the endogenous variables, *Y*, in levels. The ME-GMM technique will yield efficiency gains if the residuals of the *M*-equations are correlated and a two-step approach is thus applied to account for potential cross-equation residual correlation[22], with the weighting matrix *W* in the second step efficient GMM being defined as

$$W = N^{-1} \sum_{i=1}^{N} Z_i^{S'} e_i e_i' Z_i^S \tag{9}$$

where the vector of first step error terms $e_i = [e_{i,1} \dots e_{i,M}]'$ is derived from a consistent equation-by-equation two-stage least square (2SLS) estimation and Z_i^S is the instrument matrix as defined above. The resulting GMM estimator is thus specified in the following manner[23]

$$\hat{\delta}_{GMM} = (S'_{ZX}(W)^{-1}S_{ZX})^{-1}S'_{ZX}(W)^{-1}S_{ZY} \tag{10}$$

where $S_{ZX} = \begin{bmatrix} \frac{1}{n}\sum_{i=1}^{n} z'_{i1}x_{i1} & & \\ & \ddots & \\ & & \frac{1}{n}\sum_{i=1}^{n} z'_{iM}x_{iM} \end{bmatrix}$ and $S_{ZY} = \begin{bmatrix} \frac{i}{n}\sum_{i=1}^{n} z_{i1}y_{i1} \\ \vdots \\ \frac{1}{n}\sum_{i=1}^{n} z_{iM}y_{iM} \end{bmatrix}$

The chosen estimation procedure, ME-GMM, will thus take the endogeneity of the included variables into account hence removing the bias introduced by any correlation between the regressors and the residuals. The econometric model thus accounts for the dynamic properties of the system and the use of internal instruments in GMM controls for the endogeneity of the variables.

[21] See the technical appendix, A5, for a detailed description.
[22] The Stata command xtabond2 developed by Roodman (2009) was used in the estimation.
[23] The estimation technique is hence in line with the method adopted by Alecke, Mitze and Untiedt (2010).

6. RESULTS

Previous studies have primarily focused on estimating the relationship between the homeownership and unemployment rate by the use of correlation matrices and static OLS models. Hence, prior to elaborating on the findings from the PVARX(1) model, simpler methods will be used to analyze if an Oswald effect can be identified in the data in order to reach comparability with previous empirical studies.

Correlations can never reveal causality and one should thus be cautious in drawing strong implications from the results displayed in Table 2. However, the results do seem to imply that there is a positive relation between the unemployment rate and the owner-occupancy share. Furthermore, the three variables associated with the externalities discussed in the literature have the expected direction of the correlation. However, it is important to stress that neither does correlation reveal causality nor does it consider dynamics or cross-sectional fixed effect or other variables that the relationship might arise from. With that being said, Table 2 indicates that there might be some bearing for the Oswald effect and the possible mechanisms at play.

Correlation matrix

VARIABLES	Owner-occupancy
Unemployment	0.212***
Net migration	-0.485***
Participation	0.511***
Vacancy	-0.030

Table 2: Correlation matrix
*** p<0.01, ** p<0.05, * p<0.1

To be able to compare these findings with other studies at the macro level a static model, employing ordinary least squares (OLS), generalized least squares (GLS) and fixed-effects (FE), is estimated. The results from the OLS and GLS[24] display a significant Oswald effect in line with the point estimates found by Isebaert, Heylen and Smolders (2013) for Belgium using instrumental techniques for the homeownership share, and by Cochrane and Poot (2007) for New Zealand accounting for spatial autocorrelation and spatially lagged unemployment rates. The point estimates imply that a one percentage point increase in the owner-occupancy share yields a 0.12 percentage points increase in the unemployment rate.

The results from the static FE model with the owner-occupancy share as the explanatory variable for each of the dependent variables given by the hypotheses, unemployment rate, migration, participation and vacancy rate, are displayed in Table 3. The estimates based on a static FE model thus indicate that the Oswald effect (H1) is supported by Swedish data, with an increased owner-occupancy share yielding a positive effect on the unemployment rate with a p-value of 0.105.

Cautiousness is called for in interpreting results from the simple static OLS, GLS or FE models due to the likely problems of endogeneity and omitted variable bias. Nevertheless, the simpler static models

[24] See appendix Table A5 for results of the OLS and GLS models.

seem to support an Oswald effect in the data analyzed. As regards the three explanatory mechanisms the results seem to be mixed depending on the estimation technique chosen. The FE model does not seem to support the third hypothesis, i.e. the displacement effect, whereas the migration and consumption hypotheses do seem to find support.

Fixed effects (FE) model				
VARIABLES	H1	H2	H3	H4
Owner-occupancy	0.128	-0.042***	-0.025	-0.018**
	(0.079)	(0.013)	(0.048)	(0.008)
Year	-0.004***	0.000*	-0.001***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
Constant	7.547***	-0.083	2.182***	-0.426***
	(0.363)	(0.058)	(0.220)	(0.036)
Observations	315	315	315	315
R-squared	0.595	0.045	0.109	0.339
Number of cross-section	21	21	21	21

Table 3: Fixed effects (FE) model
Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Up to now, the possible endogeneity and dynamic structure of the variables have not been considered in the estimated model. Henceforth, the GMM procedure will be used to control for the endogeneity of the variables and the PVARX model, accounting for the dynamic structure of the system, is estimated. The model also includes the exogenous control variables, which have not been considered so far.

The GMM method relies crucially on the validity of the instrument matrix and it is thus important to properly test the validity as well as to control for possible autocorrelation in the residuals since neglected residual serial correlation can lead to biased estimates and standard deviations.

In a GMM context, a chi-square test can be used to test the over-identifying restrictions if more moment conditions than parameters are included in the model to be estimated. The test statistic, denoted the Hansen *J* statistic, has a null of a correct model specification and valid over-identifying restrictions, i.e. validity of the instruments. However, as noted by Roodman (2009), if the instrument matrix becomes too large, the test looses in power and might fail to identify invalidity in the model. In recognizing this feature of the Hansen *J* statistic as well as the importance of valid instruments in a multiple equation setting, the estimation was conducted using collapsed instrument sets[25]. Furthermore, in investigating the validity of the instrument set, a non-exhaustive instrument set is employed excluding the furthest lags of the variables until the instrument set is rendered as valid. The

[25] The collapsed technique reduces the amount of moment conditions. As stressed by Roodman (2009): "Because in the standard, uncollapsed form each instrumenting variable generates one column for each time period and lag available to that time period, the number of instruments is quadratic in *T*." Hence, as *T* increases so does the number of moment conditions and the collapsed technique is thus considered to be valuable in reducing the amount of moment conditions. For a detailed description of the collapsed technique available for the xtabond2 command in Stata, see Roodman (2009).

collapsed technique and the reduction of instrument lags reduce the number of orthogonally conditions needed in the estimation.

Another important post estimation test to conduct is the Arellano-Bond (1991) test for autocorrelation in the residuals. The test has a null of no autocorrelation and is applied to the differenced residuals. The a priori expectation is a rejection of the null for the AR (1) process in first differences due to $\varepsilon_{i,t-1}$ being present in both of the equations below

$$\Delta\varepsilon_{it} = \varepsilon_{it} - \varepsilon_{i,t-1}$$

$$\Delta\varepsilon_{i,t-1} = \varepsilon_{i,t-1} - \varepsilon_{i,t-2}$$

However, the autocorrelation test for AR(2), and higher orders, in first differences will detect if there exists any autocorrelation in the levels of the residuals.

In sum, the precise lag length included in the instrument matrix was lastly determined by the Hansen *J* statistic and the Arellano-Bond (1991) test for autocorrelation suggesting an instrument matrix consisting of lag 2 to 5 of the variables yielding 135 moment conditions for the 45 parameters to be estimated (i.e. *k*=90). The *J* statistic and AR statistic, together with the respective *p*-values, are given in Table 4 indicating that for Hansen's *J* statistic the null of validity of the instrument matrix cannot be rejected and for the Arellano-Bond (1991) test for autocorrelation the null of zero autocorrelation cannot be rejected. Furthermore, as stressed by Roodman (2009) a "perfect" *p*-value of 1.00 gives an indication of a failure of identifying invalidity in the model due to too many instruments included and as seen in Table 4, the *p*-value is both far from the conventional levels, e.g. 1, 5 or 10 %, and the worrisome "perfect" *p*-value of 1.00.

Moreover, SGMM assumes stationarity of the system. Roodman (2009) suggests that the assumption of a steady state, i.e. stationarity of the system, can be a further control for the validity of the instruments. Hence, the point estimates of the lagged left-hand side variables should be less than unity, otherwise the SGMM can be rendered as invalid. From Table 4 it is evident that all of the lagged left-hand side variables have coefficients below unity. The assumption of stationarity thus seems satisfied and these results are in line with the previously reported findings of the ADF test as shown in Table A4.

Finally, the *F*-test of joint significance indicates that the null of the right-hand side variable's coefficients being jointly equal to zero can by rejected at the conventional significance levels. Hence, the employed statistical tests jointly imply that the estimated model satisfies the key assumptions of the SGMM model and the models thus seem to be properly estimated and, as such, the estimates appear trustworthy. Lastly, dummy variables are included in the model in line with Roodman (2009) who stress the importance of including time dummies in order to prevent the most likely form of cross-sectional correlation, namely contemporaneous correlation.

Having established the validity of the estimated model, the focus will henceforth be on analyzing the estimated results, displayed in Table 4. It is evident that the point estimate indicates an Oswald effect by its sign whereas the significance level rejects its presence. In comparison with the previously estimated models, it is apparent that the effect has diminished and turned insignificant with a *p*-value of 0.63. Hence, hypothesis 1 does not seem to be supported. Moreover, as regards the explanatory factors behind the externality hypotheses, neither of the mechanisms seem to find support in the data.

As regards other potentially interesting point estimates; the unemployment rate negatively affects the owner-occupancy rate in line with prior empirics and an increased share of tertiary educated in the regional population decreases the unemployment rate and increases the vacancy rate reflecting the increasingly importance of higher education on the demand side of the labor market. The participation rate is suggested to negatively impact the unemployment rate, which is not in line with the assumptions underlying the displacement effect where increased participation is said to amplify the unemployment rate.

Dynamic model (PVARX(1))

LHS variables	RHS variables	GMM
Own_t	Own_t-1	0.854*** (0.056)
Own_t	Utot_t-1	-0.143*** (0.048)
Own_t	Vac_t-1	-1.355** (0.581)
Own_t	Net_t-1	-0.047 (0.222)
Own_t	Part_t-1	0.092** (0.043)
Own_t	Hc_t-1	-0.045* (0.026)
Own_t	Pop>55_t-1	0.004 (0.074)
Utot_t	Own_t-1	0.041 (0.084)
Utot_t	Utot_t-1	0.687*** (0.065)
Utot_t	Vac_t-1	-0.061 (0.631)
Utot_t	Net_t-1	1.060* (0.624)
Utot_t	Part_t-1	-0.544*** (0.104)
Utot_t	Hc_t-1	-0.117*** (0.038)
Utot_t	Pop>55_t-1	0.031 (0.092)
Vac_t	Own_t-1	0.002 (0.022)
Vac_t	Utot_t-1	0.054*** (0.015)
Vac_t	Vac_t-1	0.567*** (0.197)
Vac_t	Net_t-1	-0.125 (0.109)
Vac_t	Part_t-1	0.003 (0.026)
Vac_t	Hc_t-1	0.028*** (0.007)
Vac_t	Pop>55_t-1	0.026 (0.016)

Net_t	Own_t-1	0.002
		(0.024)
Net_t	Utot_t-1	-0.009
		(0.024)
Net_t	Vac_t-1	0.113
		(0.198)
Net_t	Net_t-1	0.750***
		(0.114)
Net_t	Part_t-1	-0.041
		(0.029)
Net_t	Hc_t-1	-0.008
		(0.013)
Net_t	Pop>55_t-1	-0.020
		(0.015)
Part_t	Own_t-1	-0.053
		(0.114)
Part_t	Utot_t-1	0.371***
		(0.076)
Part_t	Vac_t-1	-0.821
		(1.087)
Part_t	Net_t-1	1.575***
		(0.425)
Part_t	Part_t-1	0.402***
		(0.112)
Part_t	Hc_t-1	-0.003
		(0.039)
Part_t	Pop>55_t-1	0.312***
		(0.104)
	Region dummies	Yes
	Time dummies	Yes
	# of observations	1470
	# of instruments	135
	F- test of joint significance	11687
	H_0: Independent variables are jointly equal to zero	(0)
	Hansen J-test of over identifying restrictions	93.93
	H_0: Model specification is correct and all over identifying restrictions (all over identified instruments) are correct (exogenous)	(0.368)
	Arellano-Bond test for autocorrelation	-0.972
	H_0: No serial correlation in residuals	(0.331)

Table 4: Generalized Method of Moments (GMM)

Standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1

Note: Windmeijer (2005) finite sample corrected errors is used since the twostep approach yields downward biased standard errors. Furthermore due to a small sample, t statistics is utilized instead of z and F joint validity test instead of Wald test statistic.

The point estimates in the dynamic model estimated by the GMM technique do account for the two most sever potential problems usually overlooked by empirical estimates of the Oswald effect, namely endogeneity of the variables and dynamic properties of the system.

However, point estimates provide the long-term expected relation between two variables whereas the response over time cannot be revealed from the estimate. The OIRF analysis, on the other hand, provides guidance towards the response of an instantaneous shock to one of the error terms and reveals how the response propagates through the system over time. Hence, it is possible to evaluate the effect of one variable due to a shock of another without the need to, as demanded in interpretation of the point-estimate, restrict the response to be ceteris paribus.

The adjustment process of the unemployment rate due to a shock to the owner-occupancy share is displayed in Figure 1 below[26]. From the graph it is evident that the initial effect of an increased share of homeowners in a region is in line with the results found at the micro level; an increased share of homeownership decreases the unemployment rate initially. The unemployment rate returns to its pre-shock value within three years and then starts to increase above its pre-shock value in line with the Oswald effect. Neither the initial decrease nor the subsequent increase is however statistically significant at the 95 % level, although the decrease almost is.

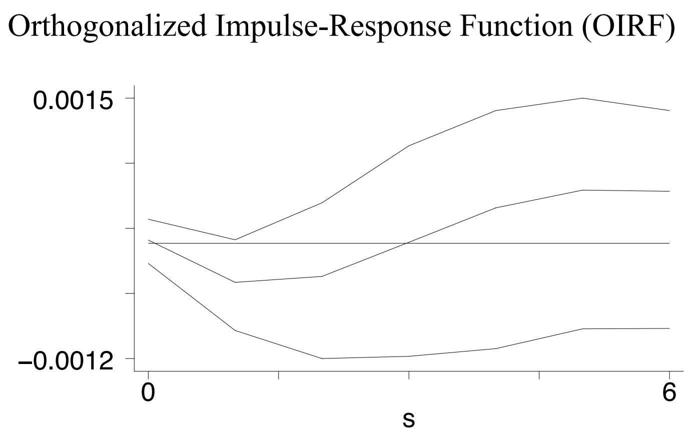

Figure 1: Orthogonalized Impulse-Response Function (OIRF) graph
Response of the unemployment rate of a shock to the owner-occupancy share
with a 95 % confidence interval based on Monte Carlo simulation with 500 reps.

The OIRF graph shows that the externality assumption of among others Blanchflower and Oswald (2013) might have some truth to it. The initial effect of a shock to the homeownership share could be interpreted as the effect found at the micro level; homeowners have a potentially better outcome in the labor market that initially decreases unemployment. The subsequent effect of returning the unemployment rate to its pre-shock value could be interpreted as the externalities produced by a large share of homeowners reversing the initial slump of the unemployment rate. However, the claim that an increased share of owner-occupancy yields a permanent increased natural rate of unemployment does not seem supported since the positive effect after about 3 years is insignificant.

[26] The OIRF results have been derived from the PVARX(1) model, including control variables, time and regional dummies.

The OIRF graph thus indicates that the Oswald effect is neither present in the short run nor significant in the long run. To shed further light on the potential presence of an Oswald effect the variance decomposition is estimated based on the OIRF results. The variance decomposition provides the proportion of the movement in a sequence due to its own shock in comparison to shocks to other variables in the system. As noted in Enders (2010), if a shock to variable *z* explains none of the variance of variable *y* at all the estimated time horizons, one can conclude that the sequence *y* is exogenous to *z*. Hence, if shocks to the owner-occupancy sequence were to explain none of the variance in the unemployment sequence then one could conclude that the Oswald effect, as suspected by among others Dietz and Haurin (2003), is non-existent and that previous positive results might be due to spurious regressions.

From the results of the variance-decomposition based on the OIRF results of the GMM model, it is evident that the homeownership share does explain some of the movement in the unemployment rate. It is however important to note the restricted magnitude; slightly below two percent of the variation in the unemployment rate is explained by the owner-occupancy share which should be compared with the participation rate explaining a substantially larger share.

Variance-Decomposition

VARIABLES	s-step ahead forecast	Own	Utot	Vac	Net	Part
Own	10	0.564	0.029	0.108	0.087	0.212
Utot	10	0.017	0.532	0.030	0.042	0.378
Vac	10	0.026	0.125	0.685	0.056	0.108
Net	10	0.030	0.065	0.049	0.787	0.068
Part	10	0.017	0.305	0.062	0.117	0.499
Own	20	0.539	0.031	0.104	0.116	0.211
Utot	20	0.017	0.531	0.031	0.044	0.377
Vac	20	0.027	0.125	0.683	0.057	0.109
Net	20	0.030	0.065	0.049	0.786	0.069
Part	20	0.017	0.306	0.062	0.117	0.498

Table 5: Variance-Decomposition

The decompositions figures are based on OIRF calculations of the GMM model

Percent of variation in row variable explained by column variable

With regards to the other OIRF graphs, displayed in Figure A1, some potentially important implications can be seen. Firstly, there does not seem to be a problem of hysteresis in the unemployment rate; it returns to its pre-shock value after about two years. Hence, the adjustment process may be sluggish but the unemployment rate is not permanently affected. Secondly, the other variables also seem to adjust towards their respective long-term levels; all of the five variables return to their pre-values after a maximum of 6 years. Thirdly, the three given externality explanations do not find support in the data; neither the vacancy rate nor the migration rate do significantly affect the unemployment rate, and the participation rate actually lowers the unemployment rate initially and begins to increase the unemployment rate after 4 years yet not significantly. Finally, from the variance decomposition it is evident that none of the supposedly endogenous variables can be rendered as exogenous since all of them do seem to explain and can be explained by the other assumed endogenous variables.

In short, the empirical assessment of the Oswald effect does neither confirm its existence nor the mechanisms supposed to cause it. To the contrary, the assessment casts serious doubt on previous results finding a positive relationship between the owner-occupancy and unemployment rate. As have been shown, the effect turns significant and appears to be of similar magnitude in relation to previous literature when employing a technique not taking into account the potential problem of endogeneity or the dynamic properties of the variables of interest. When these distortions are accounted for the effect turns insignificant.

7. DISCUSSION

In a housing market context, one could argue that Sweden is somewhat of a special case. It has been stressed by several politicians and researchers, in the public press and in studies that the regulated rentals market in Sweden hampers the development of a well-functioning rental market (Boverket 2013; Lindbeck 2013). Under such circumstances, the acquisition of a rental contract could approximate ownership. If so, the characteristics of the Swedish housing market might be different from otherwise comparable developed countries.

However, to the extent that the characteristics of the Swedish housing market make Sweden a special case in international comparisons it should, in accordance with the Oswald hypothesis, result in higher unemployment rates in Sweden than in otherwise comparable countries. This is not the situation; Sweden has had a lower unemployment rate on average over the period of interest as compared to the EU28 and in essence an equivalent unemployment rate as compared to the OECD average (OECD statistics, retrieved 11/27/2013).

Furthermore, one could expect that recent changes in Swedish housing taxation, resulting in increased transaction costs of selling one's home, in line with the Oswald hypothesis, should have strengthen a potential positive relationship between the housing market and the unemployment rate. The present government reformed the housing taxation in 2007 (prop. 2006/07:19). The capital gains taxation was increased from 20 to 22 percentage points, the deferral of a potential capital gain was constrained and an interest on the deferral was introduced (Wahlin 2013).

The reform thus increased the transaction costs associated with selling one's home and moving between regions. The reform could thus be expected to accentuate a potential positive relation between the homeownership share and the unemployment rate turning the effect significant. It would be of potential interest for further research to investigate the effects of the reform on the Oswald effect. However, such an effect, if any, does most likely have a significant time lag, which prohibits an investigation of its impact in this thesis.

Moreover, one could argue for another potential limitation, namely the cross-sectional division chosen as elaborated on in the *Data* section. It is plausible that the size of the observed regions is too large to reveal an Oswald effect with high and low homeownership counties cancelling each other out. Hence, the regional level chosen may negatively impact the estimated effect.

However, the simpler econometrical techniques imply that an Oswald effect is present and has in essence the same magnitude as the effect in international studies. Furthermore, Jonsson and Lindh (2012) find large similarities in correlation and OLS results of the Oswald effect between different aggregations of Swedish regions[27]. Hence, it does not seem as if the country characteristics and the cross-sectional level of the chosen data are as different compared to data used in previous studies and it might thus be feasible to draw general conclusions regarding the rejection of a positive relationship between the unemployment rate and the owner-occupancy share.

In further research, however, it would be of interest to examine the Oswald effect using altered aggregations of Swedish regional data and to compare the difference, if any, in outcome. Moreover, in line with the research of Holmes, Otero & Panagiotidis (2013) it would be of potential relevance to explore the dynamic response of spatial chocks by the means of GIRF analysis, in addition to the OIRF analysis presented in this thesis. Lastly, relaxing the restriction of cross-sectional homogeneity in the analysis would be of importance since such a constraint might be too strong given the nature of the data.

8. CONCLUSION

The results obtained by accounting for endogeneity and dynamic effects, in addition to conventionally included controls, reveal that the relation between a high share of homeownership and a high unemployment rate previously suggested in international research by for instance Oswald (1996, 1997a, 1997b, 1999) and Blanchflower and Oswald (2013), does not exist. Tracing the effects over time of a positive one standard deviation shock to the owner-occupancy share implies that initially the unemployment rate slumps, in line with previous micro studies, whereas after 3 years it returns to its pre-shock value. These results are in stark contrast to the hypothesized Oswald effect where a high share of homeowners is supposed to generate externalities to the regions yielding increased unemployment rates.

The OIRF analysis presented in this thesis implies that the unemployment rate initially slumps and the results are thus in apparent contrast to the Oswald effect. However, the externality effects stressed by among others Blanchflower and Oswald (2013) and Laamanen (2013) could explain the restoration of the equilibrium unemployment rate, i.e. the natural rate of unemployment. The reason for why correlation matrices, OLS and FE models pick up the effect might thus be the longer time path of the restoration compared to the initial effect. However, it should be stressed that the hypothesis of a permanently increased natural rate of unemployment due to an increased homeownership share is not supported by the results obtained in this thesis but instead rendered as false. Hence, at least in the Swedish context, the argument that the share of owner-occupancy is an important determinant for the unemployment rate can thus be dismissed. Therefore, from the results at hand, it is evident that the adverse relationship is non-existent and thus, due to the positive social effects of owning, it may very well be rational to politically promote owning at the expense of renting.

[27] Jonsson and Lindh (2012) examine the relationship between the unemployment rate and owner-occupancy share be the means of correlation matrices and OLS regressions using Swedish data spanning from 1992 to 2001. The analysis is carried out on NUTS2 level regions, NUTS3, LA and municipalities and the overall result of the OLS regressions is an insignificant relationship between the two variables. However, the effect is larger the larger is the cross-sectional level chosen.

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APPENDIX

A1 – Variables

VARIABLES	Code	Content	Unit	Source
Owner-Occupied	own	Homeowners and tenant-ownership	Share of population 20-64 years	Own calculation based on Statistics Sweden
Total unemployment rate	utot	Total unemployment, i.e. including open unemployment and labor market programs	Share of population 20-64 years	Statistics Sweden
Net-migration	net	Interregional net-migration	Share of population 20-64 years	Own calculation based on Statistics Sweden
Participation rate	part	Participation rate	Share of population 20-64 years	Own calculation based on Statistics Sweden
Vacancies	vac	Remaining available positions in the labor market at the end of the period	Share of population 20-64 years	The Swedish Unemployment Agency and Statistics Sweden
Human capital	hc	Tertiary schooling	Share of population 20-64 years	Statistics Sweden
Population > 55 years	pop55	The amount of the regional population that is 55 years or above	Share of the total population	Own calculation based on Statistics Sweden

Table A1: Variables utilized in the analysis

Owner-occupancy

The housing variable consists of three forms (homeowner, tenant-owner and rental), which for the purpose of this study was aggregated into two divisions, owner (homeowner and tenant-owner) and rental, with the owner-occupancy share being calculated using the following formula

$$own_{it} = \frac{own\_house_{it} + own\_tenant_{it}}{population_{it}}$$

Migration

The migration rate employed in the analysis is defined as the number of migrants in the working age population relative to the working age population as a whole, as indicated below with *i* denoting region and *t* time period. The rate have been calculated based on Statistic Sweden's regional in- and out-migration figures according to the following equation

$$net_{it} = \frac{in\ migration_{it} - out\ migration_{it}}{population_{it}}$$

Participation

The participation rate has been calculated based on Statistic Sweden's unemployment and employment figures according to the following equation

$$part_{it} = \frac{utot_{it} + employment_{it}}{population_{it}}$$

Vacancy rate

The vacancy data was gathered from the Swedish Unemployment Agency and was originally in quarterly form. I hence calculated the annual data as a mean of the quarterly data in line with the Agency's method for computing annual national level data. Furthermore, the vacancy data is available in two forms, new and remaining available vacancies. I chose to employ the remaining available vacancies in the analysis, in line with previous research of the Swedish labor markets (SOU 2007:35).

$$vac_{it} = \left(\frac{1}{4}\sum_{j=1}^{4} vacancy\_quarterly_{j,i,t}\right) \Big/ population_{it}$$

Descriptive statistics

VARIABLES		Mean	Std. Dev.	Min	Max	Observations
Owner-Occupancy	Overall	.7054	.0503	.5330	.7960	N = 315
	Between		.0504	.5907	.7911	n = 21
	Within		.0099	.6477	.7637	T = 15
Total unemployment rate	Overall	.1100	.0272	.0560	.2090	N = 315
	Between		.0177	.0814	.1441	n = 21
	Within		.0210	.0659	.1749	T = 15
Net migration	Overall	-.0025	.0040	-.0123	.0100	N = 315
	Between		.0034	-.0083	.0049	n = 21
	Within		.0022	-.0088	.0046	T = 15
Participation rate	Overall	.8681	.0202	.794	.912	N = 315
	Between		.0187	.8255	.8951	n = 21
	Within		.0086	.8366	.8914	T = 15
Vacancies	Overall	.0051	.0020	.0017	.0173	N = 315
	Between		.0011	.0037	.0074	n = 21
	Within		.0016	.0006	.0155	T = 15
Human capital	Overall	.2934	.0541	.1935	.4550	N = 315
	Between		.0452	.2418	.4033	n = 21
	Within		.0313	.2271	.3499	T = 15
Population > 55 years	Overall	.3152	.0277	.2342	.3703	N = 315
	Between		.0229	.2582	.3424	n = 21
	Within		.0163	.2723	.3577	T = 15

Table A2: Descriptive statistics

Note: Overall statistics are ordinary statistics that are based on 315 observations. “Between” statistics are calculated on the basis of summary statistics of 21 NUTS-3 regions (entities) regardless of time period, while “within” statistics is calculated based on each of the 15 time periods regardless of region.

A2 – Regional divisions

REGIONS
01 Stockholms län
03 Uppsala län
04 Södermanlands län
05 Östergötlands län
06 Jönköpings län
07 Kronobergs län
08 Kalmar län
09 Gotlands län
10 Blekinge län
12 Skåne län
13 Hallands län
14 Västra Götalands län
17 Värmlands län
18 Örebro län
19 Västmanlands län
20 Dalarnas län
21 Gävleborgs län
22 Västernorrlands län
23 Jämtlands län
24 Västerbottens län
25 Norrbottens län

Table A3: Cross-sectional division
The regional level utilized in the analysis is NUTS 3
Note: From 1.1.2007, Heby municipality, with about 13 000 inhabitants,
is included in 03 Uppsala County and excluded from 19 Västmanlands County

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A3 – Stationarity

Results from Augmented Dickey Fuller (ADF) test

VARIABLES	Regions	P-value < 0.20	Coefficient < 0
Net-migration	21	20	21
Total unemployment rate	21	21	21
Participation rate	21	21	21
Vacancies	21	18	21
Human capital	21	21	21
Population > 55 years	21	20	21
Owner-Occupied	21	15	20[a]

Table A4: Augmented Dickey Fuller (ADF) test preformed separately on each of the regional series

a. Dalarnas Län has a coefficient > 0

Note: The ADF test has a null of non-stationarity ($Coefficient \in \{0, -2\}$) and an alternative hypothesis of stationarity ($Coefficient > 0$). The ADF test was preformed including a drift component and zero lags.

A4 – Empirical results

Ordinary Least Squares (OLS) and Generalized Least Squares (GLS)

VARIABLES	H1 OLS	H1 GLS	H2 OLS	H2 GLS	H3 OLS	H3 GLS	H4 OLS	H4 GLS
Own	0.120***	0.124**	-0.039***	-0.040***	0.206***	0.053	-0.001	-0.005
	(0.024)	(0.055)	(0.004)	(0.009)	(0.019)	(0.040)	(0.002)	(0.004)
Year	-0.004***	-0.004***	0.000	0.000*	-0.001***	-0.001***	0.000***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Constant	7.550***	7.549***	-0.084	-0.083	2.110***	2.158***	-0.432***	-0.430***
	(0.551)	(0.362)	(0.092)	(0.058)	(0.448)	(0.223)	(0.045)	(0.036)
Observations	315	315	315	315	315	315	315	315
R-squared	0.402		0.239		0.283		0.236	
# of groups		21		21		21		21

Table A5: Ordinary Least Squares (OLS) and Generalized Least Squares (GLS)
Column variables as dependent variables
Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

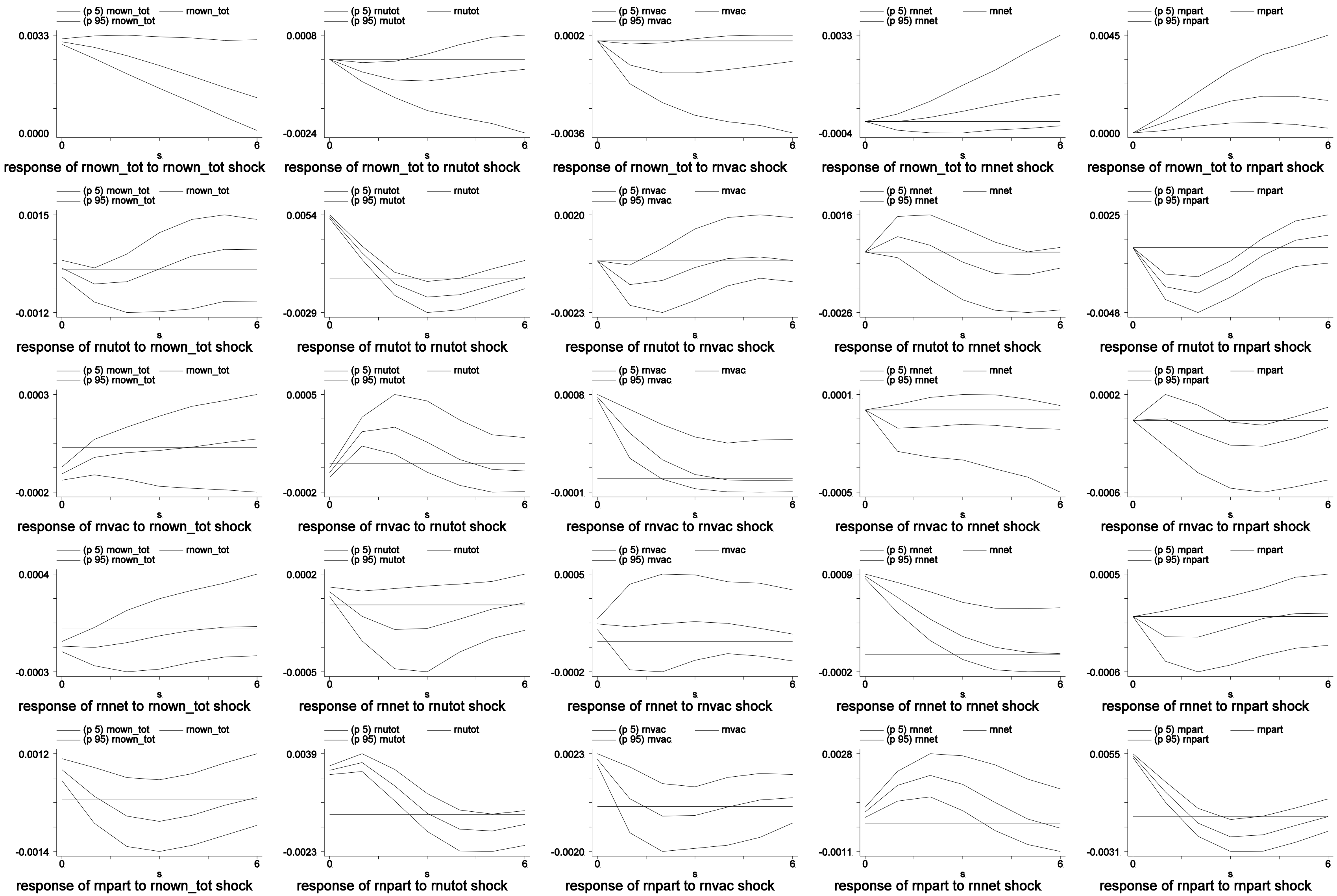

Figure A1: Orthogonalized Impulse-Response Functions (OIRF) graphs

The OIRFs are accompanied with a 95 % confidence interval based on Monte Carlo simulation with 500 reps.

Note: The confidence intervals are displayed with faulty legends due to a minor error in the OIRF code. The correct name would be the respondent variable and not the shock variable, as it is displayed now.

A5 – Technical Appendix

(Panel) Vector Autoregressive Regression

The following section is based on articles by Mitze and Stephan (2013) and Canova and Ciccarelli (2013).

In this thesis the system of equations in a dynamic model was estimated utilizing vector autoregressive, VAR, techniques in a panel setting. In a VAR model all variables are treated as endogenous and interdependent, both in a dynamic and static sense. The VAR(1) model is expressed as

$$Y_t = A_0 + A_1 Y_{t-1} + u_t \tag{A1}$$

where Y_t is a M vector of the endogenous variables, A_1 is a polynomial in the lag operator and contains all dynamic coefficients. The error terms are white noise (WN) with $E(u_t) = 0$ and $E(u_t u_t') = \Sigma_u$. A_0 consists of all deterministic components of the data: constants, seasonal dummies, time dummies, regional dummies etc.

A general variation of (A1) is to allow for a set of exogenous variables resulting in a VARX (1) model

$$Y_t = A_0 + A_1 Y_{t-1} + B_1 X_{t-1} + u_t \tag{A2}$$

where B_1 is a coefficient matrix of the exogenous variables.

Panel VARs (PVAR) have the same structure as do VAR models and assumes that all the included variables are endogenous and interdependent but with the inclusion of a cross-sectional dimension. PVARs were first employed by Holtz-Eakin et al (1988) and a PVAR(1) have the following notation

$$Y_{it} = A_{i0} + A_{i1} Y_{i,t-1} + u_{it} \tag{A3}$$

where Y_{it} is a vector of M variables for each unit $i = 1, 2, \dots, N$ where i could indicate region, firm etc. With M variables, u_{it} is a $M \times 1$ vector of random disturbances. In the setting of this thesis, the error terms are assumed to have the typical structure including fixed effects and a reminder error term such that $u_{it} = u_i + v_t + e_{it}$. Contemporaneous correlations are controlled for by inclusion of year dummies and regional effects by regional dummies.

Furthermore, when a panel VARX (PVARX) model is consider the representation for a PVARX(1) is as follows

$$Y_{it} = A_{i0} + A_{i1} Y_{i,t-1} + B_{i1} X_{i,t-1} + u_{it} \tag{A4}$$

The PVAR, and by the same token the PVARX, model thus have three characteristic features. Firstly, lags of all endogenous variables of all units enter the model for unit i, usually denoted dynamic interdependencies. Secondly, u_{it} are generally correlated across i, denoted static interdependencies among cross sections (regions, etc.). Lastly, the intercept, slope and variance of shocks to u_{it} may be unit specific, usually denoted cross sectional heterogeneity.

For the purpose of this study however the third set of features were restricted such that cross-sectional homogeneity is assumed, i.e. the slope and variance of shocks to u_{it} is assumed to be the same across $i = 1,2, \dots, N$. The resulting model is hence represented as

$$Y_{it} = A_0 + A_1 Y_{i,t-1} + B_1 X_{i,t-1} + u_t \qquad \text{(A5)}$$

Such that

$$Y_{it} = [own_{it},\ utot_{it},\ net_{it},\ part_{it},\ vac_{it}]'$$
$$X_{it} = [pop55_{it},\ hc_{it}]'$$
$$A_0 = [year, region]$$
$$i \in \{1,2, \dots ,21\}$$
$$t \in \{1,2, \dots ,15\}$$

Generalized Method of Moments Estimators

The following section is based on the econometrical textbook by Hayashi (2000).

In this thesis the joint GMM estimation of more than one simultaneous equation was utilized, denoted multiple-equation GMM (ME-GMM).

The system of $m \in \{1,2, \dots , M\}$ linear equations for $i \in \{1,2, \dots , n\}$ observations[28] can be written in the following manner:

$$y_{1i} = X'_{1i}\delta_1 + u_{1i}$$
$$\vdots$$
$$y_{Mi} = X'_{Mi}\delta_M + u_{Mi}$$

Where y_{mi} is the dependent variable of the *m*-th equation for observation i and X_{mi} is the k_m-dimensional vector of regressors. Let Z_{mi} be the vector of l_m instruments for the *m*-th equation such that

$$E(Z_{mi}u_{mi}) = 0 \ \ \forall\, m \in \{1,2, \dots , M\}$$

Writing the system in matrix notation yields

$$y_1 = X_1\delta_1 + u_1$$
$$\vdots$$
$$y_M = X_M\delta_M + u_M$$

where X_m contain the *n* observations of the right-hand side variables in the *m*-th equation such that

$$X_m = \begin{pmatrix} X'_{m1} \\ \vdots \\ X'_{mn} \end{pmatrix}$$

[28] For the purpose of this thesis: *M*=5 and *N*=(15-1)*21=294.

In a similar manner, y_m contain the n observations of the left-hand side variables in the m-th equation such that

$$y_m = \begin{pmatrix} y_{m1} \\ \vdots \\ y_{mn} \end{pmatrix}$$

and equivalently, u_m

$$u_m = \begin{pmatrix} u_{m1} \\ \vdots \\ u_{mn} \end{pmatrix}$$

Furthermore, define the following

$$X = \begin{pmatrix} X_1 & \dots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \dots & X_M \end{pmatrix}$$

$$Y = \begin{pmatrix} y_1 \\ \vdots \\ y_M \end{pmatrix}$$

$$U = \begin{pmatrix} u_1 \\ \vdots \\ u_M \end{pmatrix}$$

$$\delta = \begin{pmatrix} \delta_1 \\ \vdots \\ \delta_M \end{pmatrix}$$

The equation system can be written in compact form by noting that X is a $(nM) \times k$ matrix where $k = k_1 + \cdots + k_M$ and Y and U are $(nM) \times 1$ vectors and δ is a $k \times 1$ vector

$$Y = X\delta + U$$

The sample moment conditions in this system to be used for estimation can be defined in the following manner, considering the Ml vector of sample correlations between the exogenous variables, Z, and errors, U, such that

$$(I_M \otimes Z)'U = (I_M \otimes Z)'(Y - X\delta)$$

where Z denotes the $n \times l$ matrix of observations on the exogenous variables. Furthermore, let W be an $M \times M$ weight matrix. The system estimator, i.e. the ME-GMM estimator, is obtained by solving

$$\min_{d \in \mathbb{R}^k} (Y - Xd)'(I_M \otimes Z)W(I_M \otimes Z)'(Y - Xd)$$

And the GMM estimator is thus given by

$$\hat{\delta} = (X'(I_M \otimes Z)W(I_M \otimes Z)'X)^{-1}X'(I_M \otimes Z)W(I_M \otimes Z)'(Y) \qquad \text{(A6)}$$

Where W is defined as

$$W = \begin{bmatrix} W_{11} & \cdots & W_{1M} \\ \vdots & \ddots & \vdots \\ W_{M1} & \cdots & W_{MM} \end{bmatrix}$$

where each element, W_{im}, is an $l \times l$ symmetric matrix[29].

The ME-GMM estimators can be written out in full by the following notation

$$\begin{bmatrix} \hat{\delta}_1 \\ \vdots \\ \hat{\delta}_M \end{bmatrix} = \begin{pmatrix} X_1' Z W_{11} Z' X_1 & \cdots & X_1' Z W_{1M} Z' X_M \\ \vdots & \ddots & \vdots \\ X_M' Z W_{M1} Z' X_1 & \cdots & X_M' Z W_{MM} Z' X_M \end{pmatrix}^{-1} \begin{pmatrix} X_1' Z W_{11} Z' Y_1 + \cdots + X_1' Z W_{1M} Z' Y_M \\ \vdots \\ X_M' Z W_{M1} Z' Y_1 + \cdots + X_M' Z W_{MM} Z' Y_M \end{pmatrix}$$

In this thesis the three-stage least square, 3SLS, estimator was utilized, which is a special case of the more general ME-GMM with the restriction of on the one hand conditional homoscedasticity and on the other hand a common set of instruments across all equations.

$$E(U_i U_i{}' | Z_i) = \sum$$

$$z_{i1} = \cdots = z_{1M} = Z_i$$

where $\sum$ is some positive definite $M \times M$ matrix. The 3SLS moment conditions is then given by the $Ml \times l$ vector $U_i \otimes Z_i$ and the 3SLS efficient weight matrix is denoted such that

$$W_{3SLS} = \begin{bmatrix} \sigma_{11} E[Z_i Z_i'] & \cdots & \sigma_{1M} E[Z_i Z_i'] \\ \vdots & \ddots & \vdots \\ \sigma_{1M} E[Z_i Z_i'] & \cdots & \sigma_{MM} E[Z_i Z_i'] \end{bmatrix} = \sum \otimes E[Z_i Z_i']$$

Thus, the sample weight matrix is defined as

$$\widehat{W}_{3SLS} = \widehat{\sum} \otimes \frac{1}{n} Z' Z$$

where $\widehat{\sum}$ is the $M \times M$ matrix of estimated error cross moments estimated using the 2SLS residuals

$$\widehat{\sum} = \begin{bmatrix} \hat{\sigma}_{11} & \cdots & \hat{\sigma}_{1M} \\ \vdots & \ddots & \vdots \\ \hat{\sigma}_{M1} & \cdots & \hat{\sigma}_{MM} \end{bmatrix} = \frac{1}{n} \sum_{i=1}^{n} \hat{u}_i \hat{u}_i'$$

Then, the following expression denotes the 3SLS estimator written by matrix notation

$$\hat{\delta}^{3SLS} = \left(X'(I_M \otimes Z)\left(\widehat{\textstyle\sum}_n^{-1} \otimes (Z'Z)^{-1}\right)(I_M \otimes Z)'X \right)^{-1} X'(I_M \otimes Z)\left(\widehat{\textstyle\sum}_n^{-1} \otimes (Z'Z)^{-1}\right)(I_M \otimes Z)'Y$$

[29] Note that the single-equation GMM estimator is a special case of the multiple-equation GMM estimator with weighting matrices $W_{im,n} = 0$ for $i \neq m$.

Which, equivalently written can be denoted by

$$\hat{\delta}^{3SLS} = \left(S'_{ZX}\widehat{W}_{3SLS}S_{ZX}\right)^{-1}S'_{ZX}\widehat{W}_{3SLS}S_{ZY} \tag{A7}$$

In the single equation GMM (SG-GMM) case, the weighting matrix is expressed as

$$\hat{\Sigma} = \begin{bmatrix} \hat{\sigma}_{11} & \cdots & 0 \\ \vdots & \ddots & \vdots \\ 0 & \cdots & \hat{\sigma}_{MM} \end{bmatrix}$$

Hence, in SG-GMM one underlying assumption is that equation *h* and *m* are not correlated resulting in $\sigma_{hm} = 0$.