SPATIAL SPILLOVER IN HOUSING CONSTRUCTION

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Abstract

A model is proposed in which building contractors have regional preferences so that construction in different regions are imperfect substitutes. Contractors are hypothesized to use housing-under-construction as a buffer between starts and completions. The model also hypothesizes spatial and national spillovers in construction. Although the government does not engage in housing construction, it is assumed to have regional preferences in the initiation of housing construction. Spatial panel data are used to test the model and investigate the determinants of housing construction in Israel. Regional heterogeneity is expressed by regional specific effects in housing starts and completions. Because the spatial panel data are nonstationary, we use spatial panel cointegration methods to estimate the model. The estimated model is used to calculate impulse responses which propagate over time and across space.

Keywords: spatial panel cointegration, housing construction, space-time impulses

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“Virtually every paper written on housing supply begins with the same sentence: While there is an extensive literature on the demand for housing, far less has been written about supply.” DiPasquale (1999)

1. Introduction

As noted by DiPasquale and many others the empirical determination of house prices has attracted much more empirical attention than the empirical determination of housing construction. This continues to be so even now. This asymmetry is puzzling because house prices vary inversely with the stock of housing (Smith 1969, DiPasquale and Wheaton 1994, Bar Nathan et al 1998). Therefore a complete account of house price behavior requires analysis of both sides of the housing market, the demand for housing and its supply.

The extant research on housing construction has been largely concerned with national housing construction (Ball et al 2010). In this paper, we focus on the determinants of regional housing construction. Our motivation stems from a variety of reasons. First, regional house prices and construction vary considerably and systematically. Therefore, national housing parameters might not be relevant to specific regions. Second, national aggregation of regional housing markets might be inappropriate. Indeed, it is possible to reject a hypothesis nationally due to aggregation bias, when the hypothesis is valid regionally. Third, since regional panel data are inevitably more informative than their national counterparts, it is easier to test hypotheses using regional panel data than national data. Fourth, national models of housing supply do a poor job in capturing the unique local and regional factors that bear upon supply. Finally, to our best knowledge there is no published research on regional housing construction.

Attention has recently been drawn to local phenomena such as topography, zoning and building regulations in the determination of housing construction (Meen and Nygaard 2011 and Saiz 2010). The price elasticity of supply of new housing is expected to vary inversely with the degree of inflexibility in zoning and land use policy as well as with topographical difficulties that raise the cost of building. Since these parameters are quintessentially local, it makes more sense to estimate local or regional models rather than national models, which ignore local heterogeneity.
Regional models are not simply national models applied regionally. This is because regional housing construction is unlikely to be independent, especially if constructors operate in more than one region. Building contractors may choose to operate in regions where profits are higher, or they may have local preferences. We distinguish between relative profitability, which establishes a relationship between regional construction through regional house prices, and spatial spillovers. The latter may be induced, for example, by scale economies in which building costs in a region are affected by construction in neighboring regions, thereby inducing a spatial lag in building construction.

We propose a simple model of regional housing markets in which people prefer to live where, everything else given, housing is cheaper and building contractors prefer to build in regions where construction is more profitable. In the model there may also be spatial spillovers in housing construction and migration. In previous work (Beenstock and Felsenstein 2010) we tested this model using spatial panel data for house prices in Israel. Our main conclusion was that models are miss-specified unless they incorporate spatial spillovers in house prices. In the present paper we turn our attention to regional housing construction, and use spatial panel data to test whether housing construction models are miss-specified if they omit spatial spillovers in housing construction.

Since these data are nonstationary we use the methodology of panel cointegration to test hypotheses regarding the determination of housing construction. Indeed, a methodological aspect of the paper concerns hypothesis testing with spatially dependent panel data that are nonstationary\(^1\).

2. Theory and Methodology

2.1 The Price Elasticity of Supply of Housing Construction

The price elasticity of supply of new housing is made up of two key components. First, if house prices increase (relative to building costs) contractors have a greater incentive to build on land that is already available for housing. Marginal plots that were previously empty will be built upon and the housing stock will increase. Also, contractors will build more intensively (high rise) if building costs vary directly with the number of floors. Furthermore, marginal housing intended for re-designation (for

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\(^1\) Studies in housing supply typically ignore nonstationarity. For an exception see Mayer and Somerville (2000a).
offices, shops etc) will be retained as housing since it is more profitable, and offices and shops will be re-designated as housing. The latter does not directly affect construction but it affects the supply of housing.

Whereas the first component takes the designation of land use to be fixed, the second component assumes that land use is endogenous. If the price of housing increases, land use will be re-designated in favor of housing, which will increase new housing construction. This applies to privately owned land and publicly owned land. However, the price elasticity might be greater when land is owned privately. If land use is entirely regulated the second component will be zero because privately owned land cannot be re-designated. Also, planning permission required to build high-rise housing will adversely affect the elasticity of supply of new housing construction. However, planning permission and zoning are unlikely to be completely independent of house prices. Expensive housing makes for political unpopularity. Therefore, the second component is unlikely to be zero.

2.2 Economics of Regional Housing Markets

Two theoretical models have informed the empirical analysis of housing construction. The first relates construction to changes in house prices and the second to the level of house prices. The former treats housing as an asset to be supplied to the market if there is disequilibrium, expressed in changes in house prices (Blackley 1999, Hwang and Quigley 2006). The latter treats the production of new housing as any other product, which forms the basis of the “stock-flow” model originally proposed by Smith (1969). This model is essentially a dynamic capital asset pricing model since the price of housing is determined in the market for housing as an asset, while the flow of this asset is determined by construction, which depends upon the level of house prices.

The basic version of this model consists of two equations. In the first, the demand for housing equals supply and this determines house prices in the short run. The second equation determines housing construction, which responds to house prices. Subsequently, the housing stock adjusts over time to its long run level (Topel and Rosen 1988). The construction industry smooths-out investment over time, and

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2 As discussed below, the government has tended to sell land for housing construction when house prices were high.
3 This model dates back to Witte (1963) and has been applied in many countries including by Smith (1969) for Canada, Kearl (1979) for the United States, and Bar Nathan et al (1998) for Israel. It also features in numerous macroeconomic texts such as Dornbusch and Fischer (1990), Sachs and Larrain (1993) and Mankiw (2003).
house building is a lengthy process protracted by institutional constraints due to planning delays. Investors are encouraged to smooth construction in developed sites with permits (Mayer and Somerville 2000b). In the stock-flow model new housing competes with the existing housing stock. Since the latter is much greater than the former the market power of constructors is greatly limited. It is for this reason that the in the stock-flow model it is assumed that constructors operate within a competitive environment.

We "spatialize" the stock-flow model of the housing market. Suppose there are two regions A and B in which the population (N) is fixed so that $N_A + N_B = N$ where $N_A$ and $N_B$ are naturally positive, and $t$ labels discrete time. The population choosing to live in A is determined through the following linear migration model:

$$N_A = \varphi_0 - \varphi_1 P_A + \varphi_2 P_B$$

(1)

where $P_A$ and $P_B$ denote house prices in regions A and B respectively. The coefficients $\varphi_1$ and $\varphi_2$ reflect regional residential preferences and imply that regions are imperfect residential substitutes for each other. If they are perfect substitutes $\varphi_1 = \varphi_2 = \infty$. At the other extreme, if there is no substitution at all $\varphi_1 = \varphi_2 = 0$. If $\varphi_0$ increases region A becomes more attractive to live in.

We assume that housing construction costs, such as cement, raw materials and labor, are the same in A and B since these inputs are tradable. Contractors choose to build where it is more profitable. However, there is imperfect substitution between building in A and B if constructors have regional preferences too, or their expertise is region-specific. Given everything else, contractors therefore build more in A if housing is more expensive in A and less expensive in B. Housing construction, denoted by $B$, is determined as follows in regions A and B:

$$B_A = \eta_{A0} + \eta_{A1} P_A - \eta_{A2} P_B$$

(2)

$$B_B = \eta_{B0} + \eta_{B1} P_B - \eta_{B2} P_A$$

(3)

where $\eta_{A0}$ and $\eta_{B0}$ express productivity in construction in regions A and B respectively. The latter may be induced by physical factors that are region specific such as weather conditions and topology (Saiz 2010). The housing stocks at the beginning of period $t$ in the two regions are defined as:

$$H_{jt} = H_{j-1} + B_{j-1} - d_{j-1}$$

(4)  \hspace{1cm} j = A, B

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4 The model is assumed to be linear for analytical convenience. $N_A$ is bounded between 0 and N.
where \( d \) denotes demolitions. The regional housing market is in equilibrium when \( N_{jt} = H_{jt} \).

We solve the model for house prices under the simplifying assumption that \( d = \delta H_{t-1} \) where \( \delta \) is a common demolition rate. House prices are dynamically and spatially correlated according to the model so that current house prices in region A are related to lagged house prices in regions A and B, as well as current house prices in region B:

\[
P_{At} = \frac{1}{\varphi_1} \left[ \varphi_0 - \eta_{A0} - \eta_{A1} P_{At-1} + \varphi_2 P_{Bt} + \eta_{A2} P_{Br-1} - (1 - \delta) H_{At-1} \right]
\]

Current house prices in region A vary inversely with the local housing stock and construction productivity \((\eta_{A0})\), and vary directly with the autonomous demand to live in A \((\phi_0)\). The solution for house prices in region B has the same form as equation (5):

\[
P_{Bt} = \frac{1}{\varphi_1} \left[ Q - \varphi_0 - \eta_{B0} - \eta_{B1} P_{Br-1} + \varphi_1 P_{At} + \eta_{B2} P_{Ar-1} - (1 - \delta) H_{Br-1} \right]
\]

Equation (5) may be used to generate the following long-term solution\(^5\) for house prices in region A:

\[
P_A = \pi_0 + \pi_1 P_B - \pi_2 H_A
\]

\[
\pi_0 = \frac{\varphi_0 - \eta_{A0}}{\varphi_1 + \eta_{A1}} \quad \pi_1 = \frac{\varphi_2 + \eta_{A2}}{\varphi_1 + \eta_{A1}} \quad \pi_2 = \frac{1 - \delta}{\varphi_1 + \eta_{A1}}
\]

Equation (7) establishes that the long-term spatial lag coefficient on house prices is \( \pi_1 \). A similar result applies to region B. Notice that \( \pi_2 \) varies inversely with the elasticity of supply of construction in region A. Notice also, that \( \pi_1 \) varies inversely with this elasticity but it varies directly with \( \eta_{A2} \).

A crucial assumption in equations (2) and (3) is that apart from substitution effects induced by \( \eta_{A2} \) and \( \eta_{B2} \) the regions are otherwise independent in terms of supply. A further source of dependence would be induced by specifying spatial lags in equations (2) and (3). Thus in equation (2) we may add \( \lambda_A B_{Bt} \), and in equation (3) we may add \( \lambda_B B_{At} \) where \( \lambda \) denotes the spatial lag coefficient. If \( \lambda > 0 \) regional construction is complementary and induces mutual crowding-in. If \( \lambda < 0 \) regional construction induces mutual crowding-out. Of course, \( \lambda_A \) and \( \lambda_B \) might have opposite signs.

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\(^5\) Obtained by setting variables at time \( t \) to equal their value at \( t-1 \).
2.3 Regional Housing Policy

In Israel housing construction is entirely undertaken by private contractors. The government does not build houses directly. Nevertheless, housing construction is a major component of the government's regional policy. The government initiates housing construction in specific regions by offering for tender building rights on land vested in the Israel Land Authority (ILA). The Ministry of Housing & Construction (MOH) encourages contractors to compete for its tenders by defraying a fraction of the development costs. In this way the government subsidizes construction in regions where it wishes to initiate construction for housing. Contractors sell MOH initiated housing in the private housing market; they do not sell it to the government. It is in this way that MOH incentivizes contractors to build in specific locations.

Given everything else, there will be more construction in regions where MOH initiates more building (denoted by G). However, such building might crowd-out private building (denoted by P). Contractors who in any case intended to build in the region might simply build MOH projects instead of private projects. On the other hand, if they are credit constrained, the financial perks in MOH contracts might enable contractors to build private housing that otherwise would have not been possible. Therefore if MOH initiates 100 housing units, total construction will increase by less than 100 if there is crowding-out and it will increase by more than 100 if there is crowding-in.

Unfortunately there are no systematic data on the subsidies embodied in MOH contracts. We assume that these subsidies vary directly with MOH-initiated housing construction. Specifically, let \( Z = \frac{G}{B} \) denote the share of MOH-initiated housing in total construction (B) in the region where \( B = G + P \). If \( \ln B = \mu Z \) it may be shown that the coefficient of crowding-out is:

\[
\frac{dP}{dG} = \frac{(\mu - 1)B - \mu}{B + \mu}
\]

If \( \mu \) is positive, MOH-initiated housing (G) increases total construction. It will crowd-out private construction if \( \mu \) is approximately less than 1 and it will crowd it in otherwise.

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6 The subsidy for each MOH tender is known, but these subsidies have not been aggregated into an index.
2.3 The Econometric Model

We use spatial panel data containing NT observations to estimate equations (2) and (3). The basic model is:

\[ \ln B_i = \alpha_i + \eta \ln (P_i / C_i) + \phi \ln (\tilde{P}_i / \tilde{C}_i) + \gamma \ln (\tilde{P}_i / C_i) + \lambda \ln \tilde{B}_i + \mu Z_i + \pi \tilde{Z}_i + u_i \]  

(9)

where C denotes building costs and tildes denote spatial lags, e.g.:

\[ \tilde{B}_i = \sum_{j=1}^{N} w_{ij} B_j \]  

(10)

where \( w_{ij} \) denote exogenous spatial weights row-summed to unity and \( w_{ii} = 0 \). The main hypotheses are that regional housing construction varies directly with profitability, hence \( \eta > 0 \), and it varies directly with MOH regional incentives, hence \( \mu > 0 \). Equation (9) includes three spatial effects. First, if profitability increases among the neighbors of region i contractors will engage in spatial substitution, hence \( \phi < 0 \). See Meen and Nygaard (2011) for an example of such a spatial lag estimated from cross-section data. Secondly, if regional incentives received by the neighbors of region i induce spatial substitution in construction \( \pi \) will be negative. However, if construction in region i and its neighbors are complementary \( \pi \) may be positive. Third, if there are positive spatial spillovers in construction \( \lambda \) will be positive.

Apart from a spatial effect on profitability (\( \phi \)), equation (9) includes a national effect (\( \gamma \)). If local and neighboring profitability are given, an increase in national profitability might affect local construction in two ways. First, substitution may take place beyond neighboring regions, which would make \( \gamma \) negative. Secondly, an increase in national profitability has a positive effect on national construction. If national and local construction are complements then \( \gamma \) may be positive.

Since (see below) all the variables that feature in equation (9) are nonstationary (but are stationary in first differences) equation (9) is panel cointegrated if the residuals (\( u \)) are stationary. If the residuals are not stationary, the parameter estimates obtained from equation (9) are spurious (Phillips and Moon 1999). If equation (9) is cointegrated the parameter estimates are super-consistent\(^7\), including estimates of \( \lambda \).

The same would apply to parameters such as \( \eta \) in the event that house prices are endogenous. Superconsistency implies, for example, that \( \text{plim}( \hat{\lambda} ) = \lambda \) despite the fact

\(^7\)This would not be the case if the data are stationary, in which event estimation should be by ML or IV (Anselin 1988).
that \( \tilde{B} \) depends on \( B \). In the stock-flow model house prices are weakly exogenous because they are determined in the asset market, and the stock of housing is large relative to the flow of new housing (Topel and Rosen 1988).

If equation (9) is cointegrated the residuals are generally autocorrelated but the roots of the autocorrelation model are less than one by definition. The residuals may also be spatially autocorrelated in which case \( u_{it} \) is correlated with \( \tilde{u}_{it} \). Spatial autocorrelation reduces efficiency but does not induce bias or inconsistency in the parameters estimates. However, more efficient estimates of the parameters may by obtained by estimating equation (9) by SUR (seemingly unrelated regression).

Unfortunately data on building costs (\( C \)) are only available nationally. This may not matter as far as materials are concerned, but it may matter for labor costs. We assume, \textit{force majeur}, that regional building costs have a national component, a fixed region specific component (\( c_i \)) and a random component (\( s_{it} \)), i.e. \( C_{it} = c_i + C_t + s_{it} \) in which case \( c_i \) is absorbed into the specific effect and \( s_{it} \) is absorbed into the residual and \( C_t \) replaces \( C_{it} \) in equation (9). If the data are stationary the latter would induce attenuation bias. This problem does not arise if the data are nonstationary. Gyourko and Saiz (2006) report that construction costs vary widely in the United States. However, in a small country, such as Israel, this issue is likely to be less important.

3. The Data

3.1 House Prices

Since the early 1970s Israel's Central Bureau of Statistics (CBS) has published house price indices for nine regions (see map). These indices are constructed from transactions data, which are also used by CBS to construct a hedonic price index for the country as a whole. The spatial panel data (1987-2010) are plotted in Figure 1. They show, as expected, that housing is systematically more expensive in the core than in the periphery and that the regional ranking of house prices has remained quite stable over time. During the 1990s immigration from the former USSR increased Israel's population by 20 percent and real house prices doubled. House prices peaked in 1999-2000 after which they fell by about 30 percent. The resurgence in house prices since 2007 largely resulted from the Bank of Israel's decision to cut interest rates following the Subprime Crisis. Since we have explored these data before (Beenstock and Felsenstein 2010) we focus on housing construction.
3.2 Housing Construction

CBS publishes data on housing starts and completions by units and square meters. In what follows we use housing starts measured in square meters. We have used these data to construct housing starts for the nine regions for which house prices are available. The result is plotted in Figure 2, which shows that with the possible exception of Krayot (near Haifa) construction has had a positive trend in all regions. Krayot has systematically had the least number of housing starts, whereas Tel Aviv tended to have the most. The “spaghetti” effect in Fig 2 results from the fact that, in contrast to house prices, the regional league table in housing construction has varied over time.

Figure 3 plots MOH-initiated housing starts in the nine regions, which fall into two distinct groups. The first comprises North, South, Center and Jerusalem where most of MOH starts have been concentrated (especially South). In the second group there has been relatively little MOH activity. This largely reflects the fact that public land reserves in these regions are low. On the whole the government has been responsive to market forces; it has sold more land when house prices are more expensive. For example, following the wave of mass immigration from the former USSR in the early 1990s house prices almost doubled, and the government released land for housing. This explains the spike in 1992 in Fig 3 (especially in the South).

3.3 Panel Unit Root Tests

Panel unit root tests for logarithms of these variables are reported in Table 1. The IPS statistic tests the null hypothesis that the data are nonstationary. The LM statistic tests the null hypothesis that the panel data are stationary. Here and in the panel cointegration test statistics reported below these test statistics have been transformed into z which has a standard normal distribution:

\[ z_k = \frac{\sqrt{N}[S_k - E(S_k)]}{sd(S_k)} \Rightarrow N(0,1) \]  

where k labels the particular statistic e.g. IPS, and E(S) and sd(S) are the expected value and standard deviation of S obtained by Monte Carlo simulation.

According to IPS one may reject the null hypothesis that the log level of construction (housing starts measured in square meters) is nonstationary since \( z_{-IPS} \) (-2.4) is smaller than its critical value of -1.96. This is surprising since Figure 1.1 shows that the mean level of construction has, on the whole, been growing over time. By
contrast Hadri’s LM test clearly rejects the hypothesis that these data are stationary, since \( z_{-}L_{-}M_{-}H \) (3.67) exceeds 1.96. Ideally these tests should be mutually consistent. However, in the case of construction they are in apparent conflict. The conflict is apparent because rejecting one null hypothesis is not logically equivalent to accepting its opposite. The same apparent conflict arises in the case of housing completions.

Since the log first differences of housing starts and completions are stationary according to LMH and IPS, we consider these housing construction data to be difference stationary. This conflict is less pronounced in the case of MOH construction since \( z_{-}L_{-}M_{-}H \) is marginally smaller than its critical value. However, we also assume that MOH construction is difference stationary, which means that \( \ln Z \sim I(1) \).

### Table 1 Panel Unit Root Tests: 1987-2010

<table>
<thead>
<tr>
<th></th>
<th>( z_{-}IPS )</th>
<th>( z_{-}LMH )</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>( d = 0 )</td>
<td>( d = 1 )</td>
</tr>
<tr>
<td>House prices</td>
<td>-1.53</td>
<td>-3.96</td>
</tr>
<tr>
<td>Housing starts</td>
<td>-2.40</td>
<td>-8.41</td>
</tr>
<tr>
<td>Completions</td>
<td>-2.61</td>
<td>-7.78</td>
</tr>
<tr>
<td>Starts (MOH)</td>
<td>-3.43</td>
<td>-5.92</td>
</tr>
<tr>
<td>Housing under</td>
<td>-1.7</td>
<td>-5.61</td>
</tr>
<tr>
<td>construction</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: \( z_{-}IPS \) is the \( z \) statistic based on Im et al (2003), and \( z_{-}LMH \) is based on Hadri (2000). Two augmentations or lag truncations are specified. Data in logarithms (except housing under construction), and \( d \) denotes the order of differencing.

A similar conflict occurs in the case of the logarithm of house prices; according to IPS we may reject the null hypothesis that house prices are nonstationary, but according to IPS we may reject the null hypotheses that they are stationary. Since both IPS and LMH concur that these data are stationary in first differences, we assume that they are difference stationary.

Table 1 also includes housing under active construction (U, also measured in 1000s of square meters). The relationship between this variable and starts (S) and completions (F) is:

\[
U_t = U_{t-1} + S_{t-1} - F_{t-1}
\]  
(12)

Since S and F are cointegrated I(1) variables, equation (12) implies that \( \Delta U \sim I(0) \) in which case \( U \sim I(1) \), as indicated in Table 1 by the LMH test statistic.
3.4 Spatial Panel Unit Root Tests

The test statistics in Table 1 assume that the units in the panel are independent. Baltagi et al (2007) report that panel unit root tests which ignore spatial autocorrelation are reasonably sized provided that the spatial autocorrelation coefficient is sufficiently small (less than 0.4). In Table 2 we report critical values for ρ for the following DGP:

\[ Y_t = \alpha_t + \rho Y_{t-1} + \theta Y_t + \epsilon_t \]

(13)

where θ induces spatial dependence in the Dickey-Fuller regression. Equation (13) is a first order AR and SAR model. When θ = 0 this is equivalent to the IPS statistic expressed in terms of ρ-bar. When ρ = 0 it is equivalent to the BFF test statistic (Beenstock, Feldman and Felsenstein, 2012) for a spatial unit root. For example if N = T = 25 the critical value of ρ-bar is 0.661 at p = 0.05. If ρ-bar exceeds this critical value, the null hypothesis of nonstationarity cannot be rejected. If the panel data are spatially dependent the critical value of ρ-bar decreases slightly in Table 2 with θ.

Table 2 shows, as expected, that the critical value of ρ-bar varies directly with T and N.

**Table 2 Critical Values for Spatial Panel Unit Roots (ρ-bar)**

<table>
<thead>
<tr>
<th>N</th>
<th>p</th>
<th>0 = 0</th>
<th>0 = 0.04</th>
<th>0 = 0.2</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>1%</td>
<td>5%</td>
<td>10%</td>
<td>1%</td>
</tr>
<tr>
<td>25</td>
<td>T=10</td>
<td>0.32784</td>
<td>0.37372</td>
<td>0.39774</td>
</tr>
<tr>
<td></td>
<td>T=25</td>
<td>0.62973</td>
<td>0.66106</td>
<td>0.67869</td>
</tr>
<tr>
<td></td>
<td>T=50</td>
<td>0.80901</td>
<td>0.82487</td>
<td>0.8328</td>
</tr>
<tr>
<td>100</td>
<td>T=10</td>
<td>0.4032</td>
<td>0.42749</td>
<td>0.440415</td>
</tr>
<tr>
<td></td>
<td>T=25</td>
<td>0.67981</td>
<td>0.69633</td>
<td>0.70417</td>
</tr>
<tr>
<td></td>
<td>T=50</td>
<td>0.83531</td>
<td>0.84227</td>
<td>0.84625</td>
</tr>
<tr>
<td>225</td>
<td>T=10</td>
<td>0.430558</td>
<td>0.447017</td>
<td>0.455387</td>
</tr>
<tr>
<td></td>
<td>T=25</td>
<td>0.723823</td>
<td>0.731716</td>
<td>0.735764</td>
</tr>
<tr>
<td></td>
<td>T=50</td>
<td>0.850728</td>
<td>0.855552</td>
<td>0.858108</td>
</tr>
</tbody>
</table>

Source: Beenstock and Felsenstein (2012). Based on 10,000 Monte Carlo simulations assuming ρ_i = 1 and θ equals its tabulated value.

Table 2 suggests that if N and T are relatively small the IPS test statistic under-rejects the null hypothesis. Therefore, the results in Table 1 are conservative as far as IPS is concerned.
4. Results

4.1 Regional Housing Starts

We estimate equation (9) with regional fixed effects and SUR. The latter allows the residuals ($u_t$) to be correlated, but not necessarily spatially correlated. Since the data are nonstationary the parameter estimates have non-standard distributions, in which case $t$ -- statistics do not indicate statistical significance unless the covariates happen to be strictly exogenous, which is not the case here. We use group cointegration test statistics (Pedroni 2004) designed for panel data, which allow for heterogeneity in the autoregressive behavior of the residuals ($u_t$). Since parameter estimates typically have non-standard distributions, $t$ - tests and other test statistics that assume the parameter estimates are normally distributed are invalid. We therefore test for statistical significance by dropping variables from the model. If this induces cointegration failure we conclude that the variable or variables concerned are statistically significant.

We follow our previous work by using a spatial weighting matrix that takes account of both relative size and distance. Hence:

$$w_{ij} = \frac{POP_i}{POP_i + POP_j} \times \frac{1}{d_{ij}}$$

Where POP denotes the mean population in the data, and $d_{ij}$ is the Euclidean distance between i and j.

Table 3 Estimates of Equation (9): Housing Starts

<table>
<thead>
<tr>
<th>Model</th>
<th>$\eta$</th>
<th>$\gamma$</th>
<th>$\mu$</th>
<th>$\phi$</th>
<th>$\lambda$</th>
<th>$\pi$</th>
<th>GADF</th>
<th>GPP</th>
<th>ECM</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.312</td>
<td>0.495</td>
<td>1.098</td>
<td>-0.594</td>
<td>0.584</td>
<td>-0.433</td>
<td>-3.457</td>
<td>-3.872</td>
<td>-3.94</td>
</tr>
<tr>
<td>2</td>
<td>0.290</td>
<td>0.639</td>
<td>-0.705</td>
<td>0.650</td>
<td>0.476</td>
<td>-3.546</td>
<td>-3.987</td>
<td>-4.96</td>
<td></td>
</tr>
<tr>
<td>3</td>
<td>0.304</td>
<td>0.470</td>
<td>0.967</td>
<td>-0.548</td>
<td>0.515</td>
<td>-3.434</td>
<td>-3.819</td>
<td>-3.83</td>
<td></td>
</tr>
<tr>
<td>4</td>
<td>0.258</td>
<td>0.668</td>
<td>-0.716</td>
<td>0.730</td>
<td>-3.576</td>
<td>-4.010</td>
<td>-5.37</td>
<td></td>
<td></td>
</tr>
<tr>
<td>5</td>
<td>0.428</td>
<td>-0.031</td>
<td>1.321</td>
<td></td>
<td></td>
<td>-3.141</td>
<td>-3.556</td>
<td>-3.05</td>
<td></td>
</tr>
</tbody>
</table>

Notes: Estimation by SUR with regional fixed effects. GADF: group (1st order) ADF panel cointegration $z$-statistic. GPP: group (1st order) Phillips-Perron panel cointegration $z$-statistic. ECM: the $t$-statistic on the estimate of $\rho$ in the error correction model: $\Delta \ln S_t = \gamma_i + \rho \Delta \ln S_{t-1} + \xi \Delta \ln \bar{S}_t + \nu_t$

Results are reported in Table 3. Model 1 specifies all the variables in equation (9) and serves as an unrestricted model. The local price elasticity of supply in model 1 is
0.312, the national price elasticity is 0.495, and the spatial price elasticity is -0.594. The latter shows that spatial substitution in construction is strong, while the former shows that national and local construction are complements. MOH incentives increase construction. The sum of these elasticities (0.213) implies that a general increase in house prices of one percent raises housing starts by 0.213 percent. The estimate of $\mu$ (1.098) means that MOH construction crowds-in private construction. The spatial lag coefficient ($\lambda$) is slightly larger than a half. Finally because $\pi$ is negative, MOH construction has a negative spatial spillover effect.

The cointegration test statistics (GADF and GPP) greatly exceed their critical values. Recall that according to equation (11) these are standard normal variables. Therefore, the p-values are close to zero. We also report an ECM test for cointegration, based on the estimated residuals of model 1. The ECM test statistic is expected to be negative\(^8\). Although these test statistics ignore spatial dependence, they are so significant that they are unlikely to be misleading. We do not report t-statistics and related tests because, as mentioned, the parameter estimates have non-standard distributions.

To determine whether a covariate is statistically significant we drop it from the model. If the result ceases to be cointegrated this indicates that the covariate in question is statistically significant. For example, if $\mu = 0$ as in model 2 GADF = -3.546 instead of -3.457 as in model 1. This suggests that $\mu$ is not significantly different from zero. Table 3 reports a number of restricted models, which indicate that the group panel cointegration test statistics are insensitive to the various restrictions tested. Model 4 omits building incentives granted by the Ministry of Housing; the cointegration test statistics hardly change, suggesting that these incentives do not significantly affect construction. Model 5 omits all the spatial variables, and the panel cointegration test statistics weaken. Nevertheless, they are still statistically significant.

Figure 4 plots the estimated residuals of model 3 in Table 3. This spaghetti graph indicates that the residuals, on the whole, mean-revert to zero. However, the residuals for Haifa are an exception, as indicated by the (1st order) ADF and PP statistics reported in Table 4. Table 4 also shows that there is widespread regional heterogeneity in these mean-reverting tendencies; it is strongest in Jerusalem and the South and it is weakest in Haifa and the North. Table 4 further shows widespread

\(^8\) Critical values are given by Westerlund (2007) under the assumption that $\zeta = 0$ and $T > 100$. 
heterogeneity in regional fixed effects. The largest fixed effect is, not surprisingly, in the North since this is the largest region, and it is smallest in Krayot since this is the smallest region.

Table 4 Regional Heterogeneity (Model 3)

<table>
<thead>
<tr>
<th>Fixed Effect</th>
<th>ADF</th>
<th>PP</th>
</tr>
</thead>
<tbody>
<tr>
<td>Jerusalem</td>
<td>0.070</td>
<td>-3.072</td>
</tr>
<tr>
<td>Haifa</td>
<td>-0.742</td>
<td>-1.022</td>
</tr>
<tr>
<td>Tel-Aviv</td>
<td>-0.645</td>
<td>-1.941</td>
</tr>
<tr>
<td>Dan</td>
<td>-0.307</td>
<td>-2.740</td>
</tr>
<tr>
<td>Center</td>
<td>0.997</td>
<td>-1.526</td>
</tr>
<tr>
<td>South</td>
<td>0.401</td>
<td>-3.025</td>
</tr>
<tr>
<td>Sharon</td>
<td>0.334</td>
<td>-3.183</td>
</tr>
<tr>
<td>North</td>
<td>1.341</td>
<td>-1.149</td>
</tr>
<tr>
<td>Krayot</td>
<td>-1.443</td>
<td>-1.804</td>
</tr>
</tbody>
</table>

4.2 Spatial Panel Cointegration Tests.

The z-GADF and z-GPP statistics reported in Table 3 use critical values which assume that the panel units are independent. We have used Monte Carlo simulation to calculate the critical values for the group-rho statistic for various values of T and N when there is one covariate (Table 5). If the data are not spatially correlated ($\theta = 0$) the critical value of the group-rho statistic is 0.5867 ($N = 25, T = 15, p = 0.05$), which increases to 0.6824 when $\theta = 0.2$. It is therefore easier to refute the null hypothesis (no cointegration) when the data are spatially dependent.

Table 5 Critical Values for Group Rho in Spatial Panel Cointegration Tests

<table>
<thead>
<tr>
<th>p</th>
<th>1%</th>
<th>5%</th>
<th>10%</th>
<th>1%</th>
<th>5%</th>
<th>10%</th>
<th>1%</th>
<th>5%</th>
<th>10%</th>
</tr>
</thead>
<tbody>
<tr>
<td>N=25</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>T=10</td>
<td>0.3625</td>
<td>0.4101</td>
<td>0.4354</td>
<td>0.3653</td>
<td>0.4148</td>
<td>0.4398</td>
<td>0.4124</td>
<td>0.4661</td>
<td>0.4950</td>
</tr>
<tr>
<td>T=15</td>
<td>0.5527</td>
<td>0.5867</td>
<td>0.6055</td>
<td>0.5557</td>
<td>0.5902</td>
<td>0.6091</td>
<td>0.6340</td>
<td>0.6824</td>
<td>0.7099</td>
</tr>
<tr>
<td>T=20</td>
<td>0.6516</td>
<td>0.6817</td>
<td>0.6964</td>
<td>0.6568</td>
<td>0.6863</td>
<td>0.7010</td>
<td>0.7655</td>
<td>0.8142</td>
<td>0.8415</td>
</tr>
<tr>
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<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>T=10</td>
<td>0.4481</td>
<td>0.4706</td>
<td>0.4827</td>
<td>0.4522</td>
<td>0.4734</td>
<td>0.4858</td>
<td>0.5341</td>
<td>0.5600</td>
<td>0.5743</td>
</tr>
<tr>
<td>T=15</td>
<td>0.6151</td>
<td>0.6321</td>
<td>0.6409</td>
<td>0.6199</td>
<td>0.6365</td>
<td>0.6454</td>
<td>0.7654</td>
<td>0.7953</td>
<td>0.8100</td>
</tr>
<tr>
<td>T=20</td>
<td>0.7025</td>
<td>0.7168</td>
<td>0.7236</td>
<td>0.7078</td>
<td>0.7230</td>
<td>0.7299</td>
<td>0.9260</td>
<td>0.9521</td>
<td>0.9680</td>
</tr>
<tr>
<td>N=225</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>T=10</td>
<td>0.4743</td>
<td>0.4898</td>
<td>0.4979</td>
<td>0.4777</td>
<td>0.4924</td>
<td>0.5003</td>
<td>0.5710</td>
<td>0.5899</td>
<td>0.6001</td>
</tr>
<tr>
<td>T=15</td>
<td>0.6345</td>
<td>0.6455</td>
<td>0.6509</td>
<td>0.6389</td>
<td>0.6504</td>
<td>0.6558</td>
<td>0.8123</td>
<td>0.8331</td>
<td>0.9999</td>
</tr>
<tr>
<td>T=20</td>
<td>0.7183</td>
<td>0.7281</td>
<td>0.7327</td>
<td>0.7256</td>
<td>0.7343</td>
<td>0.7395</td>
<td>0.9764</td>
<td>0.9946</td>
<td>0.9999</td>
</tr>
</tbody>
</table>

Source: Beenstock and Felsenstein (2012). Based on 10,000 Monte Carlo simulations.
4.3 Housing Completions

We "spatialize" the multiple cointegration model between starts and completions suggested by Bar Nathan et al (1998), which ensures that starts are eventually completed. The basic hypothesis is that completions ($F$) vary directly with building under construction ($U$) and starts ($S$). Contractors use buildings under construction as a buffer which lengthens when business is bad and shortens when business is good. This means that contractors slow down completion rates when business is slack and accelerate them when business is favorable. Since regional completion rates may have a spatial dimension our basic specification for completions is:

$$F_t = \delta U_t + \zeta S_t + \omega \tilde{F}_{t-1} + \zeta \tilde{U}_{t-1} + \nu \tilde{S}_{t-1} + \omega \tilde{W}_t$$

Since all the variables in equation (14) are I(1), panel cointegration requires that $\omega \sim I(0)$. If completion rates increase when construction is more profitable, $P_{it}/C_{it}$ may be specified in equation (14). However, this effect may already be captured by starts. Notice that there is no intercept term in equation (14) because $F$ must equal zero when $S = U = 0$.

<table>
<thead>
<tr>
<th>Model</th>
<th>$\delta$</th>
<th>$\theta$</th>
<th>$\phi$</th>
<th>$\zeta$</th>
<th>$\nu$</th>
<th>GADF</th>
<th>GPP</th>
<th>ECM</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.432</td>
<td>0.169</td>
<td>0.504</td>
<td>-0.074</td>
<td>-0.226</td>
<td>-4.79</td>
<td>-5.16</td>
<td>-9.28</td>
</tr>
<tr>
<td>2</td>
<td>0.432</td>
<td>0.168</td>
<td></td>
<td></td>
<td>-4.73</td>
<td>-5.12</td>
<td>-10.9</td>
<td></td>
</tr>
<tr>
<td>3</td>
<td>0.401</td>
<td>0.276$^a$</td>
<td></td>
<td></td>
<td></td>
<td>-4.59</td>
<td>-4.99</td>
<td></td>
</tr>
</tbody>
</table>

Notes: Estimation by SUR. GADF: group (1st order) ADF panel cointegration z-statistic. GPP: group (1st order) Phillips-Perron panel cointegration z-statistic. ECM: the t-statistic on the estimate of $\rho$ in the error correction model:

$$\Delta F_{it} = \rho \hat{w}_{it-1} + \zeta \Delta F_{it-1} + \zeta \Delta \tilde{F}_{it} + v_{it} \text{.}$$

Note $a$: private housing starts.

Model 1 in Table 6 is an unrestricted model with spatial spillovers. It states that contractors complete annually 43 percent of outstanding buildings under construction, and that current completions vary directly with starts. For every 10 square meters of starts there is an additional 1.7 square meters of completions. The spatial lag coefficient is 0.504, implying that completions increase with completions in neighboring regions. There are negative spatial spillovers from buildings under construction and starts, implying that contractors substitute completions between regions. The cointegration test statistics are highly significant. Indeed, their p-values are even smaller than their counterparts in Table 3.
Model 2 shows that dropping the spatial variables makes no difference to the cointegration test statistics. Therefore, these spatial variables are not statistically significant. By contrast, in Table 3 dropping spatial variables raised the p-values of the cointegration tests. We also carried out some further tests. For example, in model 2 completions vary directly with local house prices, suggesting that contractors accelerate completions when building is more profitable. However, the cointegration test statistics do not change. Model 3 is identical to model 2 except it used private housing starts rather than total housing starts. The effect of private housing starts on completions is greater than total starts, however, there is a slight deterioration in the panel cointegration test statistics.

Table 7 The Distribution of Completions

<table>
<thead>
<tr>
<th>Year</th>
<th>0</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Completions</td>
<td>16.8</td>
<td>35.9</td>
<td>20.4</td>
<td>11.6</td>
<td>6.6</td>
<td>4.0</td>
</tr>
<tr>
<td>Completion Rate</td>
<td>16.8</td>
<td>52.7</td>
<td>73.1</td>
<td>84.7</td>
<td>91.3</td>
<td>95.3</td>
</tr>
</tbody>
</table>

The completion lag implied by model 2 is represented in Table 7. It follows a cohort of 100 additional starts, which occur in year 0. What matters is not the completion of these particular houses, but the completion of housing as a whole when contractors use housing under construction as a buffer. Notice that these starts induce contractors to complete housing under construction more rapidly. Therefore completions increase in year 0. Even if it takes time to build, the effects of starts on completions is instantaneous. Completions peak in year 1 by which the completions rate is 52.7 percent. Subsequently, the completion rate increases towards 100 percent. The mean lag is 2.7 years.

4.4 Model Properties

To illustrate the properties of the multiple cointegration housing construction model we use model 4 for housing starts from Table 3 and model 2 for housing completions from Table 6. There are spatial spillovers in the former but not in the latter. The choice is made for reasons of parsimony and the p-values of the panel cointegration tests. Notice that the starts model is in logarithms but the completions model is not. Therefore the model is nonlinear. The model is completed by using equation (12) to relate building under construction to starts and completions.

We set up a base-run by carrying out a full dynamic simulation (FDS) of the model over 1988 – 2010 in which the state variables, such as house prices and MOH
starts, assume their values as in the data. Because the model contains levels of variables and their logarithms the model is nonlinear and its solutions are base dependent. We calculate impulse responses by perturbing the state variables and by comparing the perturbed FDS to the base run. In doing so, we distinguish between local, spatial and global perturbations. Due to the presence of spatial effects in the housing starts model, the impulse responses propagate over space as well as time.

The model is dynamic because of the lag between starts and completions. Since the equations for starts and completions refer to their nonstationary components, and do not embody short-term dynamics, the model refers to trend, or equilibrium behavior. A complete dynamic account would have to include error correction models for starts and completions. In the absence of error correction, the simulated impulse responses therefore refer to equilibrium responses, and their dynamics are entirely induced by the lag between completions and starts.

Table 8 Model Simulations: Housing Starts

<table>
<thead>
<tr>
<th></th>
<th>Tel Aviv</th>
<th>Jerusalem</th>
<th>Haifa</th>
<th>Center</th>
<th>Dan</th>
<th>Sharon</th>
<th>Krayot</th>
<th>North</th>
<th>South</th>
</tr>
</thead>
<tbody>
<tr>
<td>A</td>
<td>0.35</td>
<td>0.30</td>
<td>0.45</td>
<td>0.19</td>
<td>0.29</td>
<td>0.67</td>
<td>0.41</td>
<td>8.94</td>
<td>0.28</td>
</tr>
<tr>
<td>B</td>
<td>2.27</td>
<td>-0.83</td>
<td>-0.66</td>
<td>-0.83</td>
<td>-1.5</td>
<td>-1.05</td>
<td>-0.68</td>
<td>-0.98</td>
<td>-0.73</td>
</tr>
<tr>
<td>C</td>
<td>-4.35</td>
<td>-3.08</td>
<td>-3.50</td>
<td>-2.74</td>
<td>-4.37</td>
<td>-3.84</td>
<td>-3.93</td>
<td>-3.77</td>
<td>-2.84</td>
</tr>
</tbody>
</table>

A: MOH housing starts increased in North by 200,000 square meters.
B: House prices in Tel Aviv increased by 10 percent
C: Building costs increased by 10 percent

In the first simulation we increase temporarily MOH housing starts in the North in 1995 by 200,000 square meters. This is an example of a local perturbation for the North. However, from the point of view of neighboring regions this is a spatial perturbation. In the interest of space, we focus on the response of housing starts (Table 8) and housing stocks (Table 9). The former lasts for one period only because the shock lasts for one period, and because the cointegrating vector for starts contains no dynamics. The latter, as mentioned, is dynamic because of the relationship between starts and completions. Table 9 reports the response of housing stocks upto 7 years after the shock.

The direct effect on housing starts in the North is 185,327 square meters (8.94%). Housing starts increase by less than 200,000 square meters because MOH starts crowd out private starts (simulation A). The rate of crowding out in the North in 1995 was 7.3 percent; a square meter of MOH starts crowds out 0.073 square meters
of private starts. Through spatial lag effects housing starts increase in other regions. There are two types of spatial lag effect. First, there are spatial spillovers from housing starts. Second, there are spatial spillovers from MOH starts. The former spatial spillovers propagate across the regions of Israel through the spatial lagged dependent variable. Spatial spillovers for housing starts are all positive and range from 0.19 percent to 0.66 percent.

Because the perturbation is assumed to be temporary, housing starts revert to their baseline solution. However, housing stocks are permanently raised, especially in the North. It takes about 4 years for the completion – starts process to dissipate after which housing stocks settle down to their new equilibrium. By implication, completions and housing-under-construction revert to their base-run solutions. By year 7 after the shock, the housing stock in the North increases by 0.43 percent, but most of this increase has already occurred within 3 years. Housing stocks gradually increase in other regions because of the spatial spillovers in starts.

<table>
<thead>
<tr>
<th>Lag</th>
<th>Tel Aviv</th>
<th>Jerusalem</th>
<th>Haifa</th>
<th>Center</th>
<th>Dan</th>
<th>Sharon</th>
<th>Krayot</th>
<th>North</th>
<th>South</th>
</tr>
</thead>
<tbody>
<tr>
<td>A</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1</td>
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<td>.005</td>
<td>.004</td>
<td>.004</td>
<td>.002</td>
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<td>-.171</td>
<td>-.106</td>
<td>-.171</td>
<td>-.080</td>
<td>-.180</td>
<td>-.125</td>
</tr>
</tbody>
</table>

Next (simulation B), we simulate a temporary increase of house price (10% in 1995) in the Tel Aviv region, which raises housing equilibrium starts by 2.27 percent in Tel Aviv. This increase comes at the expense of housing starts elsewhere. This happens because there is spatial substitution in housing construction; there is less incentive to build outside Tel Aviv. However, this effect is mitigated by the spatial lag in housing starts. The decreases in housing starts elsewhere range from 0.68 percent to 1.5 percent. Not surprisingly, these decreases are strongest in the vicinity of Tel Aviv.
Aviv, especially Dan and Sharon. The spatial spillovers are large relative to their counterparts in the previous simulation (A). As in simulation A, it takes about 5 years for housing starts to find their way into the housing stock.

Finally (simulation C), we simulate a temporary increase in national construction costs in 1995. This is an example of a global perturbation. National construction costs affect starts in three ways. First, since local construction costs depend on national construction costs, local relative profitability in construction decreases, which adversely affects local construction in all regions. Second, if construction profitability decreases in neighboring regions, this increases local construction through the spatial lag coefficient. Third, construction profitability decreases nationally, which adversely affects local construction since local and national construction are complementary. The first and third effects are negative and the second effect is positive. However, the combined effect is negative as may be clearly seen in the simulation.

The adverse effects of construction costs on housing starts range from 2.74 percent in the Center and 4.34 percent in Tel Aviv. This heterogeneity stems from the spatial lag structure of the model, and because the spatial weights are asymmetric and vary. The spatial weights take account of relative size and distance. Therefore, the spatial effect of e.g. Tel Aviv on Jerusalem does not equal the effect of Jerusalem on Tel Aviv, and the effect of Jerusalem on Haifa differs from the effect of Jerusalem on Tel Aviv. As in simulations A and B it takes about 5 years for the housing stocks to adjust.

5. Conclusion
Using recent methodological advances in the econometric analysis of nonstationary spatial panel data and spatial panel data for Israel we have investigated the determinants of regional housing construction. We show that although housing starts vary directly with profitability as measured by house prices relative to building costs, they vary inversely with profitability in neighboring regions, i.e. there is substantial spatial substitution in housing construction. The local price elasticity of supply is about 0.3, whereas the spatial elasticity is about -0.6. This substitution effect suggests that contractors have local building preferences since they regard neighboring regions as close substitutes but not more distant regions.
Whereas neighboring regions are substitutes, we find that local and national construction are complements. If national profitability increases, this raises local construction, as well as national construction. The local elasticity of supply with respect national house prices is about 0.5. The overall price elasticity of supply is about 0.25, i.e. a general increase in house prices of 10 percent raises construction across the country as a whole by about 2½ percent.

Apart from the spatial substitution effect mentioned above, a further spatial effect is captured by the spatial lagged dependent variable in the model for housing starts. The estimated spatial lag coefficient implies that the local elasticity of construction with respect to construction in neighboring regions is about 0.6, suggesting that local construction and neighboring construction are complementary. We reconcile this complementarity and the substitution effect as follows. Contractors may regard neighboring regions as substitutes, but there are favorable synergies in regional construction. The cost of building in a region varies inversely with construction in its neighbors due, for example, to cost sharing in the use of capital equipment as well as perhaps in the use of labor. These spatial effects emphasize the difference between spatial and national modeling of housing supply. A regional model is not simply a national model applied regionally.

In Israel the Ministry of Housing and Construction does not directly engage in housing construction. Instead, it auctions off land for house building at preferential terms. We show that such building tends to crowd-in housing construction. The financial perks that accompany these auctions help constructors engage in other housing construction, suggesting that constructors are capital constrained. Therefore, housing construction initiated by MOH does not tend to crowd-out other housing construction. However, there is a spatial effect insofar as auctions in neighboring regions reduce local construction. Contractors will build less in a locality if MOH is initiating housing construction among its neighbors. This result is consistent with our finding that local and neighboring construction are substitutes.

We show that the lag between completions and starts varies inversely with the number of starts. This is consistent with the hypothesis that contractors use building under-construction as a buffer to smooth construction. They slow down the completion rate when business is quiet and increase it when business picks-up. Unlike in the case of housing starts, we find little in the way of spatial spillovers in housing completions. However, there may be a spatial lag in housing under-construction so
that local completions vary directly with housing under-construction in neighboring regions. This effect is consistent with our previous finding that local and neighboring starts are complementary.

We use the model to simulate impulse responses across space and over time. Region specific shocks propagate at three levels. They propagate over time within regions. They propagate between regions. Finally, they propagate between regions over time. We report impulse responses for MOH initiated housing, house prices and building costs. In doing so, we distinguish between local and nation-wide shocks. The reported impulse responses express the richness of the spatial specification of the model.

Finally, we draw attention to a number of econometric issues. Since the panel data are nonstationary we have used panel cointegration to test hypotheses about housing construction. It is assumed in standard panel unit root and panel cointegration tests that the units in the panel are independent. This assumption is naturally violated in spatial panel data. We have carried out Monte Carlo simulations of the sensitivity of these tests to spatial dependence between panel units. These simulations show that provided the spatial dependence is not too pronounced the critical values for standard panel unit root and cointegration tests are reasonably reliable.
Fig 1 Regional Housing Data 1987-2010
Fig 1.1 Housing Starts (1000's m²)

Figure 2 Public Sector Housing Starts (1000's m²)
Fig 3 House Prices (Set 1991’s CPI=100)

Figure 4: Residuals of Model 3
Regional Map of Israel

Regions:
1. Jerusalem
2. Tel Aviv
3. Haifa
4. Haifa Bay
5. Gush Dan
6. Sharon
7. Center
8. North
9. South
References


